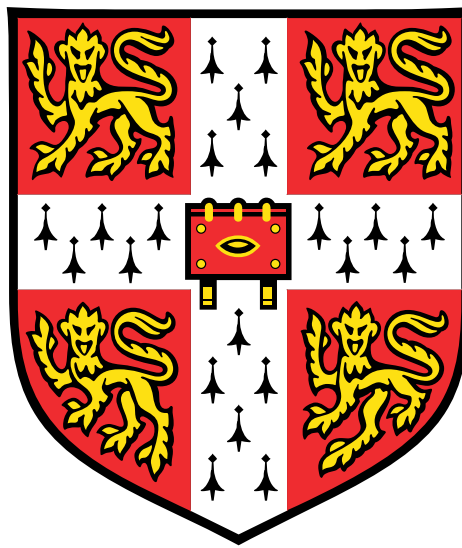


An Analysis of Monetary Policy Transmission Through Bond Yields



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This dissertation is submitted for the degree of
Doctor of Philosophy

To my loving wife.

Declaration

This dissertation is the result of my own work and includes nothing which is the outcome of work done in collaboration. It is not substantially the same as any that I have submitted, or, is being concurrently submitted for a degree or diploma or other qualification at the University of Cambridge or any other University or similar institution. I further state that no substantial part of my dissertation has already been submitted, or, is being concurrently submitted for any such degree, diploma or other qualification at the University of Cambridge or any other University or similar institution. It does not exceed the prescribed word limit of 60,000 words.

Simon Lloyd
September 2017

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Abstract

In this thesis, I study the transmission of monetary policy through the term structure of interest rates. This is an important topic because, with short-term nominal interest rates in many advanced economies close to their effective lower bound since 2008-2009, central banks have used ‘unconventional’ monetary policies, such as large-scale asset purchases and forward guidance, to stimulate macroeconomic activity by, *inter alia*, placing downward pressure on longer-term interest rates. I focus on the mechanisms through which monetary policy influences bond yields, domestically and globally, with reference to a canonical decomposition of longer-term interest rates into expectations of future short-term interest rates, and term premia.

After an introduction in chapter 1, chapter 2, *Overnight Indexed Swap Market-Based Measures of Monetary Policy Expectations*, appraises the use of overnight indexed swap (OIS) rates as measures of expected future monetary policy. Unlike federal funds futures (FFFs), which have regularly been used to construct measures of US interest rate expectations, OIS rates are available in many countries. I find that US OIS rates provide measures of interest rate expectations that are as good as those from FFFs, and that US, UK, Eurozone and Japanese OIS rates up to a 2-year horizon tend to accurately measure interest rate expectations, providing comparable cross-country measures of monetary policy expectations.

In chapter 3, *Estimating Nominal Interest Rate Expectations: Overnight Indexed Swaps and the Term Structure*, I propose a novel method for estimating interest rate expectations and term premia at short and long-term horizons: a no-arbitrage Gaussian affine dynamic term structure model (GADTSM) augmented with OIS rates. Using 3 to 24-month OIS rates, the OIS-augmented model generates estimates of the expected path of short-term interest rates out to a 10-year horizon that closely correspond to those implied by FFFs rates and survey expectations, outperforming existing GADTSMs.

I study the transmission of US unconventional monetary policies in chapter 4, *Unconventional Monetary Policy and the Interest Rate Channel: Signalling and Portfolio Rebalancing*. Using the OIS-augmented GADTSM, I carry out an event study to demonstrate that US unconventional monetary policy announcements between November 2008 and April 2013 did significantly reduce US longer-term interest rates by affecting expectations and term premia. As a result of these declines, unconventional monetary policies aided US real economic outcomes. Using a structural vector autoregression, I show that changes in interest rate expectations, linked to monetary policy signalling, had more expansionary effects on US real economic outcomes than changes in term premia, associated with portfolio rebalancing.

Chapter 5, *Long-Term Interest Rates, International Risk Sharing and Global Macroeconomic Spillovers*, assesses the international transmission of monetary policy through the term structure of interest rates between advanced economies. I present a micro-founded, two-country model with endogenous portfolio choice amongst country-specific short and long-term bonds, and eq-

uity. Within the model, US monetary policy has sizeable effects on longer-term interest rates in other advanced economies, which are similar to empirical estimates. Using the OIS-augmented GADTSM in an event study, I show that US monetary policy has led to changes in interest rate expectations in other advanced economies that amplify global spillovers, which have been partly mitigated by changes in term premia through portfolio rebalancing.

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Chapter 1

Introduction

The 2007-2008 financial crisis had profound effects on the conduct of monetary policy in advanced economies. In response to interbank market turmoil, which began in August 2007, central banks swiftly injected liquidity into these markets in order to contain volatility. However, the failure of Lehman Brothers in September 2008 led to an acute crisis in global financial markets. Liquidity support was no longer deemed sufficient, and central banks looked, *inter alia*, to interest rate policy to support financial markets and the macroeconomy. Central banks in advanced economies sharply reduced their traditional policy instrument, the short-term nominal interest rate, to historically low levels. By December 2008, the Federal Reserve's headline policy rate, the federal funds rate, was targeted at 0 to 0.25%, while UK Bank Rate reached 0.5% in March 2009.

Despite reducing short-term policy rates to their effective lower bound (ELB), central banks deemed that further stimulus was necessary to support the macroeconomy, and turned to 'unconventional' monetary policy tools, defined here as instruments beyond the traditional policy rate. These unconventional monetary policies, such as large-scale asset purchases (or quantitative easing) and forward guidance, were, *inter alia*, designed to place downward pressure on longer-term interest rates and, in turn, stimulate macroeconomic activity. Although these policies were considered part of an appropriate monetary policy response at the ELB (e.g. Clouse, Henderson, Orphanides, Small, and Tinsley, 2003; Eggertsson and Woodford, 2003; Bernanke, Reinhart, and Sack, 2004), there was limited historical precedent for their efficacy. For this reason, unconventional monetary policies drew attention to monetary policy's influence on the term structure of interest rates.

This thesis examines the transmission of monetary policy through the term structure of interest rates, providing answers to a number of important questions.

First, through what mechanisms does monetary policy influence longer-term interest rates? This is especially important when enacting policy that is directly aimed at reducing longer-term interest rates, which I answer with recourse to a canonical decomposition of longer-term interest rates into expectations of future short-term interest rates and term premia. Movements in long-term interest rates that emanate from changes in expected future short-term interest rates are associated with revisions to investors' expectations of the future monetary policy stance and a signalling channel. In contrast, changes in term premia reflect the compensation investors demand for taking on risks, and can, in part, reflect portfolio rebalancing. This decomposition is particularly salient for the study of monetary policy, because the two components have differing policy implications. If monetary policy predominantly reduces longer-term interest rates

through signalling, monetary policymakers may influence the efficacy of their policy by clearly communicating information about future policy through, for example, forward guidance. In contrast, if monetary policy predominantly affects term premia, then this motivates policies that can influence investors' portfolio positions, such as large-scale asset purchases. This decomposition of longer-term interest rates is a central strand connecting the chapters of this thesis.

Second, given this decomposition of longer-term interest rates, how can we measure these sub-components in order to empirically gauge the relative importance of the signalling versus portfolio rebalancing channels of monetary policy? I address the measurement of interest rate expectations and term premia in chapters 2 and 3.

Third, given empirical estimates of the longer-term interest rate decomposition, through what mechanisms did unconventional monetary policies influence domestic macroeconomic activity? I use the proposals of chapters 2 and 3 to empirically gauge the relative importance of signalling and portfolio rebalancing channels of monetary policy in chapters 4 and 5.

Fourth, how do monetary policy shocks globally transmit through longer-term interest rates? This is the focus of chapter 5.

In chapter 2, I investigate the use of high frequency financial market data for measuring expectations of the future policy rate at short to medium-term horizons. Historically, researchers and financial market practitioners have used federal funds futures (FFFs) rates to measure US interest rate expectations. Many authors have used FFFs rates to study the transmission of US monetary policy, and proposed methods reliant on FFFs that are widely used by monetary economists (e.g. Kuttner, 2001; Gürkaynak, Sack, and Swanson, 2005a,b; Gertler and Karadi, 2015). However, FFFs are only traded in the US, and very few similar instruments are traded in other jurisdictions. Moreover, studies have shown that FFFs, with horizons in excess of 1-year, may contain risk premia that limit their use as a measure of interest rate expectations (Piazzesi and Swanson, 2008). I propose the use of overnight indexed swap (OIS) rates as financial market-based measures of interest rate expectations that are comparable across countries. First, I compare US OIS rates to FFFs rates, and find that 1 to 11-month OIS rates provide measures of investors' interest rate expectations that are as good as those implied by FFFs rates spanning the same horizon. Second, I find that, on average, 1 to 24-month US, UK, Eurozone and Japanese OIS rates accurately measure expectations of future short-term interest rates at daily frequencies. Motivated by these results, researchers can look to OIS rates as a globally-comparable measure of shorter-term monetary policy expectations that enables them to apply many empirical methods used by monetary economists to a wider set of countries.

Chapter 3 builds on this by proposing a method for estimating interest rate expectations and term premia at short and long-term horizons. To do this, I present a no-arbitrage dynamic term structure model. These models have regularly been used to estimate interest rate expectations and term premia, but daily frequency estimates of existing model variants fail to accurately capture the evolution of interest rate expectations implied by surveys and financial market instruments (e.g. Bauer, Rudebusch, and Wu, 2012; Kim and Orphanides, 2012; Guimarães, 2014). I propose the augmentation of no-arbitrage Gaussian affine dynamic term structure models (GADTSMs) with OIS rates in order to better estimate the evolution of interest rate expectations and term premia across the whole term structure. The results in chapter 2 provide evidence that the 3 to 24-month OIS rates that I augment the model with have statistically insignificant excess returns, and so provide valid information with which to identify

interest rate expectations. The OIS-augmented model, estimated for the US, generates estimates of the expected path of short-term interest rates that closely correspond to those implied by FFFs rates and survey expectations out to a 10-year horizon, and accurately depict their daily frequency evolution. Against these metrics, the interest rate expectation estimates from the OIS-augmented model is superior to estimates from existing GADTSMs.

In chapter 4, I study the transmission of the US unconventional monetary policies enacted by the Federal Reserve from late-2008, large-scale asset purchases and forward guidance. These policies transmit to the real economy, *inter alia*, via an interest rate channel, with two sub-channels: expected future short-term interest rates, labelled signalling, and term premia, associated with portfolio rebalancing. I apply the OIS-augmented GADTSM from chapter 3 to identify these two sub-channels. Using an event study, I demonstrate that US unconventional monetary policy announcements between November 2008 and April 2013 did exert significant signalling and portfolio balance effects on financial markets, reducing longer-term interest rates. Signalling effects were particularly powerful at horizons in excess of 2-years. As a result of these declines, unconventional monetary policy aided real economic outcomes. Using a structural vector autoregression, I show that the signalling channel exerted a more powerful influence on US industrial production and consumer prices than portfolio rebalancing. In terms of long-term bond yield and industrial production effects, the signalling channel is associated with around two-thirds to three-quarters of the total effects attributed to the two channels.

I study the global transmission of US monetary policy shocks through long-term interest rates in chapter 5, where I present a micro-founded, two-country model with endogenous portfolio choice amongst country-specific equity, short-term bonds, and long-term bonds. The model provides novel insights about the different roles played by short and long-term bonds in international risk sharing and the global transmission of monetary policy shocks. Short-term bonds predominantly hedge real exchange rate fluctuations that occur immediately after a macroeconomic shock, while long-term bonds mainly hedge expected inflation and expected real exchange rate movements. The model predicts that a surprise tightening of conventional US monetary policy leads to an immediate increase in long-term interest rates in the US and other advanced economies, which align with empirical results from an event study, indicating that US monetary policy has sizeable spillover effects through the yield curve, corroborating with recent evidence in Gilchrist, Yue, and Zakrajsek (2016). I extend the empirical results by applying the OIS-augmented GADTSM from chapter 3. I find that the spillover effects of US monetary policy are closely associated with changes in interest rate expectations: a surprise tightening of US monetary policy immediately increases investors' expectations of future short-term interest rates in the US and other advanced economies. However, term premia fall in response to the same shock, attenuating global spillover effects.

This thesis is related to a long line of research studying how monetary policy works. Early studies of monetary policy transmission often relied on empirical vector autoregressions (VARs), with quarterly or monthly frequency macroeconomic data (e.g. Bernanke and Blinder, 1992; Christiano, Eichenbaum, and Evans, 1996), or the dynamic New Keynesian setup formalised in Woodford (2003). With increasingly abundant financial data, academic research now incorporates high frequency financial information to identify monetary policy shocks and better understand the transmission of monetary policy. Kuttner (2001) proposed the use of FFFs rates as measures of monetary policy surprises, with subsequent research by Bernanke and Kuttner

(2005) and Gürkaynak et al. (2005a) building on this. Gürkaynak et al. (2005a,b) and Hanson and Stein (2015) have used these high frequency techniques to document how the US yield curve reacts to monetary policy surprises in the US. Researchers have also taken steps to simultaneously benefit from high frequency financial data and lower frequency macroeconomic data. Bernanke, Boivin, and Elias (2005) is a widely cited example, but more recently Gertler and Karadi (2015) have proposed the use of high frequency changes in FFFs rates to identify monetary policy surprises within a VAR by building on contributions by Stock and Watson (2012) and Mertens and Ravn (2013).

The empirical contributions in chapters 2 and 3 of this thesis have important implications for this area of the empirical monetary economics literature. In chapter 2, I show that OIS rates, out to 2-year horizon, provide high frequency measures of interest rate expectations that are comparable across countries. This is important in view of the recent growth of a literature assessing the international transmission of monetary policy (Rey, 2016), which will benefit from financial market measures of interest rate expectations that are available outside the US and are globally comparable. Moreover, as chapters 4 and 5 indicate, the OIS-augmented decomposition of longer-term interest rates proposed in chapter 3 can be used to study the relative importance of different sub-channels of monetary policy, with important implications for the conduct of policy.

Chapter 5 of this thesis also contributes to a growing literature on the global spillover effects of US monetary policy. Rey (2014) and Passari and Rey (2015) present evidence of a global financial cycle in gross cross-border asset flows, asset prices and leverage, motivating research by Miranda-Agrippino and Rey (2015), Dedola, Rivolta, and Stracca (2016) and Rey (2016) documenting the effects of US monetary policy on the global financial cycle. Faust, Rogers, Wang, and Wright (2007), Beechey and Wright (2009), Ehrmann, Fratzscher, and Rigobon (2011) and Gilchrist et al. (2016) analyse the response of international financial markets to US monetary policy surprises. Chapter 5 extends this literature in two ways. First, by presenting a model of endogenous portfolio choice among country-specific assets — namely equity, short and long-term bonds — and, second, by applying the OIS-augmented decomposition from chapter 3 to study the specific, policy-relevant, transmission channels of international spillovers.

The policy implications of this thesis are especially timely. As central banks in advanced economies seek to raise interest rates from their ELB, my conclusions indicate that clear communication of future intentions will be of first-order importance for the efficacy of policy. Moreover, the evolution of term premia, in part due to portfolio rebalancing, will have significant implications for the international spillover effects due to monetary policy. By proposing methods for the accurate measurement of interest rate expectations and term premia, this thesis also provides tools for tracking the efficacy of monetary policy in real time over the coming years and beyond.

Chapter 2

Overnight Indexed Swap Market-Based Measures of Monetary Policy Expectations

2.1 Introduction

Researchers, policymakers and financial market participants closely monitor the evolution of expectations about the future path of monetary policy. This has been particularly apparent in recent years, as central banks have considered raising policy rates from their effective lower bound (ELB) (Lao and Mirza, 2015). Because of the keen interest in monetary policy expectations, empirical measures of investors' expectations of future short-term interest rates are highly sought-after. Within the academic literature, such measures have formed an important part of the empirical toolkit for monetary economists, informing numerous methodological contributions (e.g. Gürkaynak, Sack, and Swanson, 2005a,b; Gertler and Karadi, 2015; Cesa-Bianchi, Thwaites, and Vicondoa, 2016).

Broadly speaking, empirical measures of investors' expectations of future short-term interest rates can be categorised into three groups:¹ (i) financial market-based, where interest rate expectations are extracted from raw financial market data, such as futures and swaps; (ii) model-based, where interest rate expectations are estimated within models that use financial market data as an input, such as dynamic term structure models;² and (iii) survey-based, for instance from surveys of professional forecasters.

Financial market-based measures are the primary focus of this chapter. To date, the principal measures amongst these are federal funds futures (FFFs) rates. In a widely cited paper, Gürkaynak, Sack, and Swanson (2007b) compare the empirical success of a number of US financial market-based measures — including FFFs and eurodollar futures — as predictors of the future monetary policy stance. They conclude that, out to a 6-month horizon, FFFs dominate all other financial market instruments in forecasting monetary policy and, at longer horizons, the predictive power of many instruments is similar. However, Gürkaynak et al. (2007b) do not compare FFFs to overnight indexed swap (OIS) rates.

¹These three categories are not mutually exclusive. See chapter 3 for a proposal that incorporates both financial market and model-based measures to estimate investors' interest rate expectations at horizons up to 10 years, and Kim and Orphanides (2012) for a proposal that combines model and survey-based measures.

²See chapter 3 and the references within.

FFFs have played an important role in forging an empirical toolkit for monetary economists, enabling the study of monetary policy and its effects. However, FFFs are US-only financial instruments and very few similar instruments are traded elsewhere. Therefore, the majority of questions to which this empirical toolkit has been applied have focused almost exclusively on the US. For example, work by Kuttner (2001) and Gürkaynak et al. (2005a) has assessed the effect of US monetary policy shocks on investors' interest rate expectations, as measured by FFFs rates. Similarly, FFFs have been used to analyse the effects of US 'unconventional' monetary policies, such as large-scale asset purchases and forward guidance, on financial markets (Swanson, 2016).³

Financial market-based measures of interest rate expectations offer certain advantages over model and survey-based measures for these applications. Most importantly, financial market data is available at intraday frequencies. In comparison, model-based measures are most-widely available at monthly or daily frequencies, while survey-based measures are (at best) available at monthly frequencies. The availability of intraday data has been critical in the aforementioned literature, permitting the identification of exogenous shocks to monetary policy uninfluenced by other economic news. Moreover, financial market-based instruments circumvent potentially contentious modelling assumptions applied to model-based measures, and the limitations of sampling from a population of individuals to create survey measures of interest rate expectations.

Most recently, a burgeoning literature, motivated by Stock and Watson (2012) and Mertens and Ravn (2013), has combined high-frequency identification techniques with structural vector autoregression methods to estimate the macroeconomic effects of monetary policy shocks. Gertler and Karadi (2015) use surprise movements in FFFs rates during 30-minute windows around monetary policy announcements from Gürkaynak et al. (2005a) as exogenous instruments to structurally identify monetary policy shocks. With this literature in its infancy, the application of financial market-based measures of interest rate expectations in academic research is set to grow. Moreover, with the growth of a parallel literature assessing the international transmission of monetary policy (Rey, 2016), there is a need to find financial market measures of interest rate expectations that are available outside the US and are globally comparable.

In this chapter, I propose and test the use of OIS rates for this purpose. Since their inception in the early 2000s, OIS contracts have grown in popularity within financial markets. An OIS contract is an over-the-counter traded derivative in which two counterparties exchange fixed and floating interest rate payments. The floating interest rate on OIS contracts is the overnight interbank rate, which provides a measure of the *de facto* monetary policy stance.⁴

OIS contracts have numerous features that make them excellent candidate measures of investors' interest rate expectations. First, there is no exchange of principal, minimising counterparty risk. Second, OIS contracts do not involve any initial cash flow, minimising liquidity risk. Third, because many OIS contracts are collateralised, credit risk is also minimised (Tabb and Grundfest, 2013). Finally, unlike many LIBOR-based instruments, OIS contracts have increased in popularity following the 2007-2008 financial crisis (Cheng, Dorji, and Lantz, 2010).

Within the US, OIS rates offer potential advantages over FFFs rates too. First, OIS rates are now available at maturities in excess of 3 years. Cheng et al. (2010) state that OIS contracts tend to be liquid out to at least the 3-year horizon. FFFs are traded at up to a 3-year horizon, but remain largely illiquid at maturities in excess of 1 year. Second, the horizon of OIS contracts

³Model-based estimates of investors' interest rate expectations have also been used to assess the transmission channels of unconventional monetary policies (e.g. Gagnon, Raskin, Remache, and Sack, 2011, and chapter 4).

⁴For example, the floating reference rate on US OIS contracts is the effective federal funds rate.

on a given day aligns with the horizon of government bond yields, whereas the horizon of FFFs contracts is a specific calendar month in the future, changing only at the beginning of a new calendar month. This permits easier comparison across financial instruments than FFFs.⁵

These promising features of OIS contracts have been recognised in studies assessing the efficacy of recent ‘unconventional’ monetary policies and their effect on expectations of future short-term interest rates (e.g. Christensen and Rudebusch, 2012; Woodford, 2012, and chapter 4). Yet, despite the growing use of OIS rates as measures of interest rate expectations, no study has formally assessed the empirical success of OIS rates for this purpose.

In this chapter, I address two questions. First, how accurate are implied interest rate expectations from US OIS rates, and how do they compare to those from FFFs? This offers a useful benchmark for comparison, as the behaviour of FFFs rates has been widely studied (e.g. Gürkaynak et al., 2007b; Piazzesi and Swanson, 2008; Hamilton, 2009). Second, how accurate are implied interest rate expectations from OIS rates in other countries — specifically the UK, Eurozone and Japan? This is important for the global application of OIS rates as a financial market-based measure of monetary policy expectations.

To compare US OIS rates and FFFs, I build on the methodology of Piazzesi and Swanson (2008) and calculate *ex post* excess returns on these instruments. To perform this comparison accurately, I design a method to ensure that the horizons of OIS and FFFs contracts are identical. I create ‘portfolios’ of FFFs contracts and compare them to US OIS rates on the penultimate business day of each month. Plots of the unconditional *ex post* excess returns on OIS contracts and comparable-maturity portfolios of FFFs strongly indicate that OIS and FFFs rates contain similar information pertaining to investors’ expectations of future short-term interest rates. I find that 1 to 11-month OIS contracts provide measures of investors’ interest rate expectations that are as good as those from comparable-horizon FFFs contracts.

I then assess the global comparability of OIS rates. I first calculate the average *ex post* excess returns on US contracts using daily data to attain a benchmark against which to compare global OIS rates. The average *ex post* returns on US OIS contracts are comparable at monthly and daily frequencies, indicating that the OIS-FFF comparison is not blurred by calendar effects. Between 2002 and 2016, 1 to 24-month US OIS contracts, on average, provide accurate measures of investors’ expectations of future short-term interest rates. I then calculate the average *ex post* excess returns on UK, Eurozone and Japanese OIS contracts at the same daily frequency. I find that, on average, 1 to 24-month OIS contracts in these jurisdictions also provide accurate measures of investors’ expectations of future short-term interest rates.

These results have important implications for the future fashioning of an empirical toolkit for monetary economists. OIS rates have a useful role to play as a globally comparable market-based measure of interest rate expectations in empirical and policy research on positive and normative economic questions from a global, non-US-centric, perspective.

The remainder of this chapter is structured as follows. Section 2.2 describes FFFs and OIS contracts. Section 2.3 presents the empirical comparison of FFFs and OIS. The global comparison of OIS contracts is in section 2.4. Section 2.5 concludes.

⁵Chapter 3 proposes a method for estimating interest rate expectations out to a 10-year horizon that combines OIS rates and zero-coupon government bond yields.

2.2 Financial Market Instruments

2.2.1 Federal Funds Futures (FFFs)

FFFs contracts were introduced by the Chicago Board of Trade (CBOT) in 1988 and are unique to US financial markets. They have a variety of maturities extending to the first 35 calendar months into the future. The contracts pay out at maturity based on the average effective federal funds rate realised in the calendar month specified in the contract. For example, the first FFF settles based on the average effective federal funds rate for the calendar month in which the contract was purchased. The second FFF settles based on the average effective federal funds rate in the calendar month subsequent to purchase, and so on.

Let $p_{t,t+n}^{FFF}$ denote the price of the FFFs contract purchased on a given day during month t that settles based on the average daily effective federal funds rate (an annualised rate) during month $t+n$ (the ‘delivery month’) $\overline{f\overline{f}r}_{t+n}$, for $n = 0, 1, \dots, 35$.⁶ The contract matures at the end of the calendar month $t+n$, with settlement occurring on the subsequent day. The contract settles at “100 minus the arithmetic average of the daily effective federal funds rate during the delivery month”.⁷ The price quote is equal to 100 minus the expectation of the average daily effective federal funds rate in the delivery month. As such, the n -month FFFs rate, $i_{t,t+n}^{FFF} = 100 - p_{t,t+n}^{FFF}$, represents market participants’ expectations of the average effective federal funds rate in the delivery month. Thus, for the buyer of the contract, the *ex post* realised (annualised) excess return equals:⁸

$$rx_{t,t+n}^{FFF} = i_{t,t+n}^{FFF} - \overline{f\overline{f}r}_{t+n} \quad (2.1)$$

where $\overline{f\overline{f}r}_{t+n}$ is the *ex post* realised average daily effective federal funds rate for month $t+n$.⁹

Under the expectations hypothesis, the FFF rate $i_{t,t+n}^{FFF}$ must equal the *ex ante* expectation of the average daily effective federal funds rate $\overline{f\overline{f}r}_{t+n}$ for the contract month $t+n$:

$$i_{t,t+n}^{FFF} = \mathbb{E}_t [\overline{f\overline{f}r}_{t+n}] \quad (2.2)$$

Thus, if the *ex post* realised excess return in (2.1) has zero mean, the *ex ante* forecasting error under the expectations hypothesis also has zero mean, and the n -month FFF can be said to provide an accurate measure of expected future short-term interest rates.

2.2.2 Overnight Indexed Swaps (OIS)

An OIS is an over-the-counter traded interest rate derivative with two participating agents who agree to exchange fixed and floating interest payments over a *notional* principal for the life of the contract. The floating leg of the contract is constructed by calculating the accrued interest payments from a strategy of investing the notional principal in an overnight reference rate and

⁶That is, $n = 1$ refers to the one-month ahead contract (FF2 on financial market platforms); $n = 2$ the two-month ahead contract (FF3), and so on.

⁷See CME Rulebook, Chapter 22, 22101: www.cmegroup.com/rulebook/CBOT/V/22/22.pdf.

⁸Piazzesi and Swanson (2008) remark that (2.1) treats FFFs contracts as forward contracts, abstracting from the fact that futures contracts are ‘marked to market’. Nevertheless, they demonstrate that the empirical difference between the precise definition of the *ex post* realised excess return on FFFs contracts, which accounts for marking to market, and (2.1) is small and does not influence their results for FFFs contracts. As in the main body of Piazzesi and Swanson (2008), I therefore use (2.1) to define *ex post* realised excess returns for simplicity.

⁹The *ex post* realised average effective federal funds rate for month $t+n$ is formally calculated as the arithmetic mean of the daily effective federal funds rate for the contract month, where the rate on non-business days is defined to be the rate that prevailed on the preceding business day.

repeating this on an overnight basis for the duration of the contract, investing principal plus interest each time. The reference rate for US OIS contracts is the effective federal funds rate, while for UK, Eurozone and Japanese contracts the reference rates are SONIA, EONIA and TONAR, respectively. The ‘OIS rate’ represents the rate on the fixed leg of the contract. For a vanilla OIS contract with a maturity of one year or less, money is only exchanged at the conclusion of the contract. Upon settlement, only the net cash flow is exchanged between the parties.¹⁰ That is, if the accrued fixed interest rate payment exceeds the floating interest payment, the agent who took on the former payments must pay the other at settlement. Importantly, there is no exchange of principal at any time for OIS contracts of all maturities.

Due to the features of the contracts, OIS rates are closely linked to investors’ expectations of future overnight interest rates over the horizon of the contract. Specifically, liquidity premia on OIS contracts should be small because there is no initial cash flow and, as an OIS contract is in zero net supply, it is unclear which party would demand a liquidity premium. Counterparty risk is small because there is no exchange of principal. Moreover, because many OIS trades are collateralised, credit risk is also minimised (Tabb and Grundfest, 2013, pp. 244-245). Finally, unlike many LIBOR-based instruments, OIS contracts have increased in popularity amongst investors following the 2007-2008 financial crisis (Cheng et al., 2010).

Let $i_{t,t+n}^{OIS}$ denote the annualised n -month OIS rate in month t , the swap’s fixed interest rate. $i_{t,t+n}^{FLT}$ is the annualised *ex post* realised (net) return from the floating leg of the same contract.

The floating leg of the contract $i_{t,t+n}^{FLT}$ is calculated by considering a strategy in which an investor borrows the swap’s notional principal x , invests in the overnight reference rate and repeats the transaction on an overnight basis, investing principal plus interest each time. Let the contract trade day be denoted t_{1-s} , where s denotes the ‘spot lag’ of the contract in days.¹¹ Suppose the n -month (N -day) contract matures on the day t_N in month $t + n$.¹² The floating leg of the contract is calculated based on the realised overnight reference rate on days t_1 to t_N .¹³ Thus, the contract settlement period is given by the days t_1, t_2, \dots, t_N . The floating overnight reference rate for the OIS contract on day t_i is denoted flt_i , for $i = 1, \dots, N$. Following market convention, the expression for the floating leg of an n -month (N -day) OIS contract, purchased on day t_{1-s} , in month t is:

$$i_{t,t+n}^{FLT} = \left(\left[\prod_{j=1}^N (1 + \gamma_j flt_j) \right] - 1 \right) \times \frac{360}{N} \quad (2.3)$$

where γ_j is the accrual factor of the form $\gamma_j = D_j/360$, where D_j is the day count between business days t_j and t_{j+1} .¹⁴ To compare this floating leg to the fixed leg $i_{t,t+n}^{OIS}$, which is reported

¹⁰For contracts with maturity in excess of one year, the net cash flow exchange occurs at the end of every year.

¹¹In the US market, OIS payment calculations begin with a two-day spot lag ($s = 2$) from the trade date, so the trade day is labelled t_{-1} . The same spot lag is included in Eurozone and Japanese OIS contracts. However, the spot lag on UK contracts is zero days, so the trade day is t_1 .

¹²That is, the contract matures on day t_N , $(s - 1) + N$ days after the trade date t_{1-s} .

¹³The floating leg is calculated using the actual month lengths, not normalised month lengths.

¹⁴For example, on a week with no public holidays, the day count D_j will be set to 1 on Monday to Thursday, 3 on Friday, and 0 on Saturday and Sunday. That is, the floating overnight reference rate on a non-business day is defined as the rate that prevailed on the preceding business day. For US and Eurozone contracts, the day count is divided by 360, while for UK and Japanese contracts it is divided by 365, as per market convention.

on an annualised basis, $i_{t,t+n}^{FLT}$ is a multiple of $360/N$ in (2.3).¹⁵

From the perspective of an agent who swaps fixed interest payments for the floating rate over the notional principal x , $(i_{t,t+n}^{OIS} - i_{t,t+n}^{FLT}) \times x$ represents the payoff of a zero-cost portfolio.¹⁶ Thus, the *ex post* realised (annualised) excess return on the n -month OIS contract purchased during month t is:

$$rx_{t,t+n}^{OIS} = i_{t,t+n}^{OIS} - i_{t,t+n}^{FLT} \quad (2.4)$$

Under the expectations hypothesis, the fixed leg of the OIS contract must equal the *ex ante* expectation of the floating leg:

$$i_{t,t+n}^{OIS} = \mathbb{E}_t [i_{t,t+n}^{FLT}] \quad (2.5)$$

Thus, if the *ex post* realised excess return in (2.4) has zero mean, the *ex ante* forecasting error under the expectations hypothesis also has zero mean, and the n -month OIS contract can be said to provide an accurate measure of expected future short-term interest rates.

2.3 A Comparison of US OIS and FFFs Rates

2.3.1 Comparing OIS and FFFs Rates

(2.1) and (2.4) provide definitions for the *ex post* realised excess returns on FFFs and OIS contracts respectively. However, a direct comparison of excess returns of the two instruments on any given day is prevented by two distinct complications which motivate my empirical strategy.

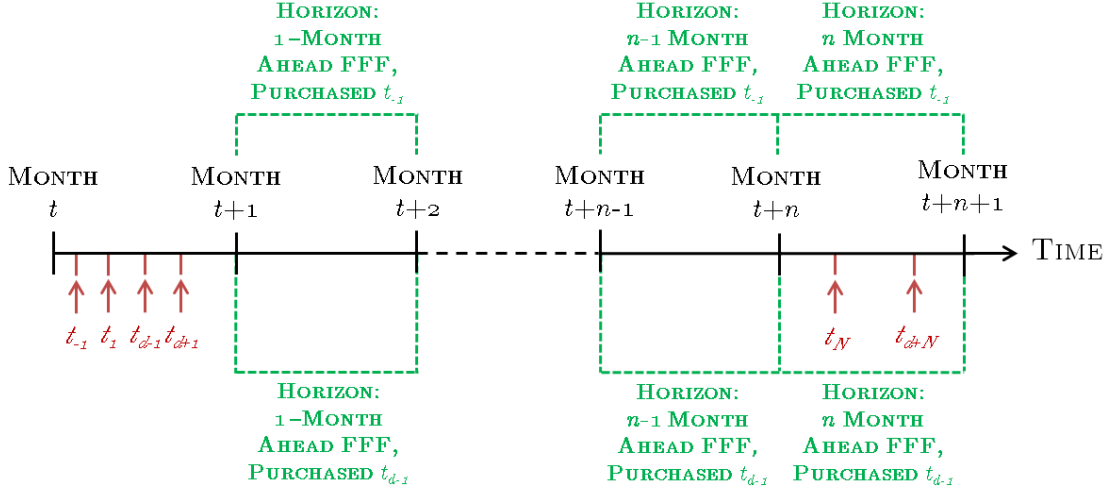
First, an n -month ahead FFF contract traded in the calendar month t has the same settlement date, regardless of the day in the month it is traded. In contrast, the n -month (N -day) OIS contract has a settlement horizon spanning the n -months (N -days) subsequent to the trade date (once the spot lag has been accounted for). This difference is depicted in figures 2.1 and 2.2. Figure 2.1 illustrates that an n -month ahead FFF contract traded on day t_{-1} in month t pertains to the same time horizon as an n -month ahead FFF contract traded on a different day t_{d-1} in the same calendar month t . Figure 2.2 demonstrates that an n -month US OIS contract, with N days to maturity and a two-day spot lag, traded on day t_{-1} settles based on the geometric average of the daily effective federal funds rate from day t_1 (accounting for the spot lag) to t_N (the maturity date), while the n -month contract traded on a different day t_{d-1} in the same month t settles based on the geometric average of the daily effective federal funds rate from day t_{d+1} to t_{d+N} .

Second, the horizon of an n -month FFF contract purchased on any day t_{-1} in month t pertains to a specific calendar month in the future $t+n$, while the horizon of an n -month OIS contract purchased on the same day t_{-1} of month t spans the n months (N days) subsequent to date t_1 (accounting for the two business day spot lag for US contracts). In figure 2.2, the n -month OIS contract traded on day t_{-1} of month t has a horizon spanning from day t_1 (accounting for the spot lag) in month t to day t_N in month $t+n$. In contrast, the n -month ahead FFF contract traded on the same day t_{-1} in month t has a horizon corresponding to the calendar month $t+n$ only (see figure 2.1).

¹⁵This accords with US and Eurozone market quoting conventions. Specifically, the fixed and floating legs of US OIS contracts are quoted according to the *Actual 360* convention. The UK and Japanese market quoting convention is *Actual 365*, so the floating rate is a multiple of $365/N$ instead (OpenGamma, 2013, p. 6).

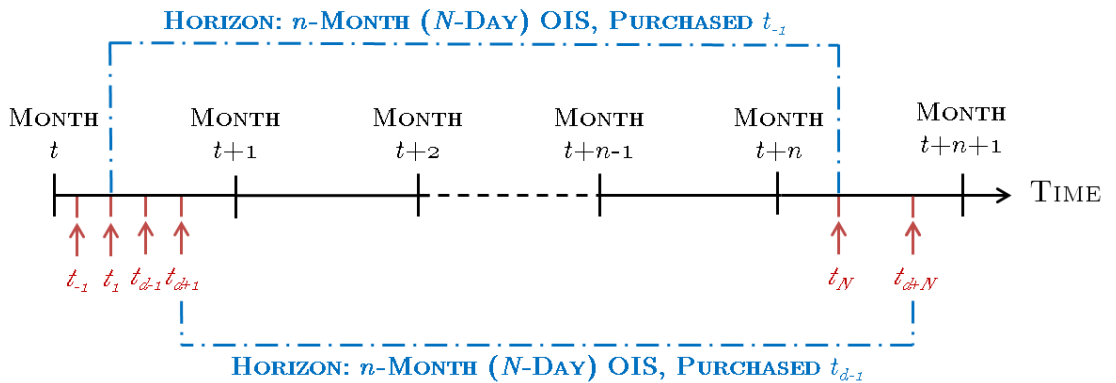
¹⁶Formally, this portfolio involves borrowing x at the floating overnight index rate at day t_1 and rolling over the borrowing to day t_N , while investing the x borrowed on day t_1 in the fixed interest rate $i_{t,t+n}^{OIS}$ to day t_N .

Figure 2.1: The Horizon of Federal Funds Futures Contracts Traded on Different Days, t_{-1} and t_{d-1} , in the Month t



Note: This figure depicts the horizons of federal funds futures (FFFs) contracts traded on different days, t_{-1} and t_{d-1} , in the calendar month t . An n -month ahead FFF contract purchased on day t_{-1} in month t pertains to the same time horizon as an n -month ahead FFF contract purchased on a different day t_{d-1} in the same calendar month t .

Figure 2.2: The Horizon of Overnight Indexed Swap Contracts Traded on Different Days, t_{-1} and t_{d-1} , in the Month t



Note: The figure depicts the horizons of US overnight indexed swap (OIS) contracts traded on different days, t_{-1} and t_{d-1} , in the calendar month t . The horizon of an n -month OIS contract purchased on day t_{-1} in month t spans from day t_1 (accounting for a two business day spot lag according to US market convention) to day t_N in month $t+n$. In contrast, the horizon of an n -month OIS contract purchased on a different day t_{d-1} of the same calendar month t spans from day t_{d+1} to day t_{d+N} in month $t+n$.

To address these complications and attain comparable measures of *ex post* realised excess returns on FFFs and OIS contracts, I perform two data transformations.

First, I construct hypothetical n -month ‘portfolios’ of FFFs contracts that are traded in month t with horizons that begin on the first day of month $t + 1$ and conclude on the final day of month $t + n$. To construct the interest rate on the n -month FFFs-portfolio $i_{t,t+n}^{FFF,port}$, I take the arithmetic average of the 1, 2, ..., n -month ahead FFFs rates on a given day in month t :

$$i_{t,t+n}^{FFF,port} = \frac{1}{N} \sum_{j=1}^n N_j i_{t,t+j}^{FFF} \quad (2.6)$$

where N_j denotes the number of days in month $t + j$ and $N \equiv \sum_{j=1}^n N_j$ is the total number of days in months $t + 1, \dots, t + n$.

This hypothetical n -month contract settles based on the arithmetic average of the daily effective federal funds rate from the first day of month $t + 1$ to the final day of month $t + n$, denoted $\overline{ffr}_{t,t+n}^{port} = \frac{1}{N} \sum_{j=1}^n N_j \overline{ffr}_{t,t+j}$.¹⁷ Thus, the *ex post* realised excess return on the hypothetical n -month portfolio of FFFs contracts, relative to the contract’s settlement, is defined as:¹⁸

$$rx_{t,t+n}^{FFF,port} = i_{t,t+n}^{FFF,port} - \overline{ffr}_{t,t+n}^{port} \quad (2.7)$$

Second, the horizon of an n -month US OIS contract will only match the horizon of an n -month portfolio of FFFs contracts on the penultimate business day of a given month because of the two-day spot lag in US OIS contracts. Figure 2.3 demonstrates this. Here, day t_{-1} is the penultimate business day of month t . The horizon of the n -month OIS contract traded on this day begins on day t_1 , the first day of the month $t + 1$, because of the spot lag in the contract. It concludes on day t_N , the final day of month $t + n$. The horizon of the hypothetical n -month portfolio of FFFs contracts, defined in (2.6), spans the same period. Because of this, I compare the *ex post* realised excess returns on n -month OIS contracts and n -month portfolios of FFFs contracts for month t on the penultimate business day of that month, using the definitions in (2.4) and (2.7) respectively.¹⁹ This yields a monthly time series of *ex post* realised excess returns for OIS and FFFs contracts that are comparable in horizon and formed on the same date.

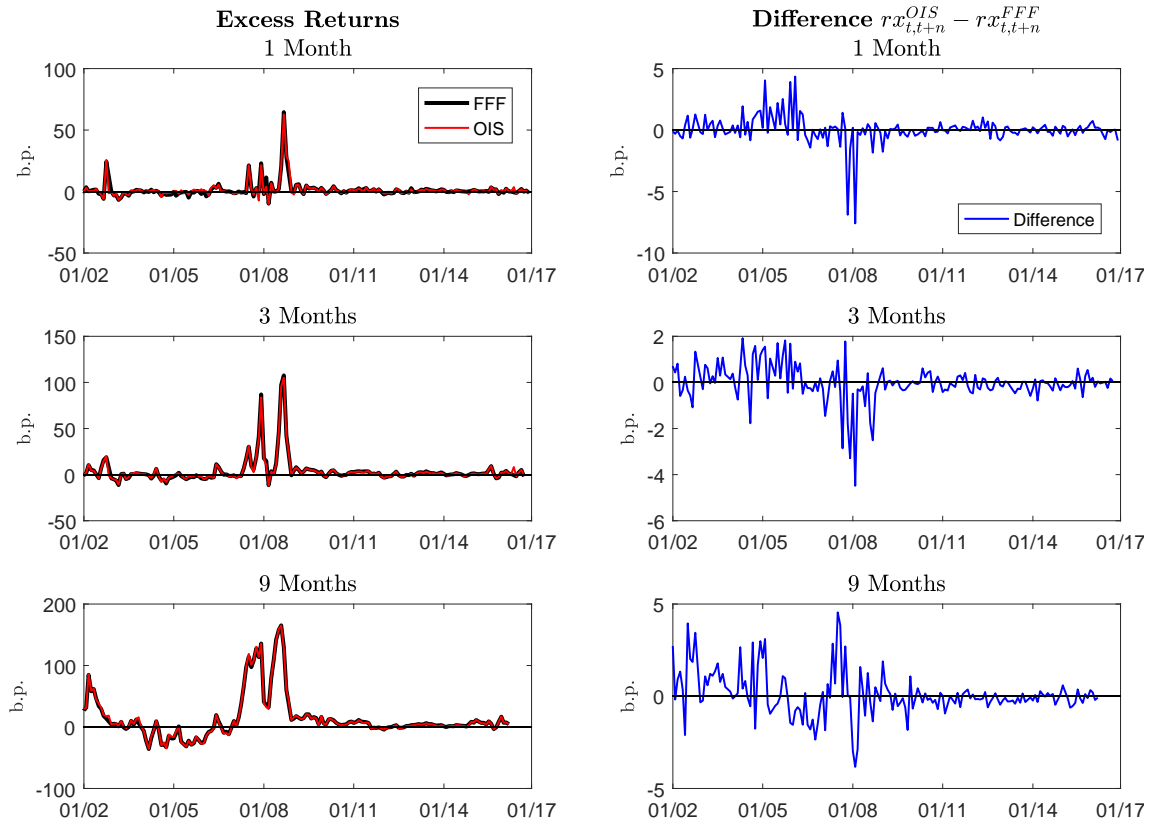
The left-hand column of figure 2.4 plots the time series of unconditional *ex post* excess returns on 1, 3 and 9-month OIS contracts and portfolios of FFFs contracts on the penultimate business day of each month. The right-hand column plots the difference between the two. These plots preview some of the conclusions from the formal empirical analysis. Notably, at all tenors,

¹⁷I use the arithmetic average in accordance with FFFs market convention. To calculate this *ex post* realised average effective federal funds rate, the rate on non-business days is defined to be the rate that prevailed on the preceding business day.

¹⁸With the exception of the 1-month horizon, the timing of receipts from an n -month OIS contract differs from an n -month portfolio of FFFs. Unlike an OIS contract, the portfolio of FFFs does not provide a single payoff at the end of its n -month horizon, but one at the end of each month as each FFFs contract matures. This is mitigated by focusing on the *ex post* realised excess returns on the portfolios, assuming that FFFs receipts prior to month $t + n$ can earn a compounded return equal to the effective federal funds rate until the portfolio matures. For example, at the end of month $t + 1$, the investor earns an excess return from the 1-month FFFs contract in the portfolio $rx_{t,t+1}^{FFF}$, which could earn compounded interest at the effective federal funds rate from the first day of month $t + 2$ to maturity at $t + n$. Because (2.7) defines an *ex post* excess return, the interest earned on $rx_{t,t+1}^{FFF}$ and the interest foregone exactly cancel, so do not feature on the right-hand side of (2.7).

¹⁹Although the horizons of $i_{t,t+n}^{OIS}$ and $i_{t,t+n}^{FFF,port}$ traded on day t_{-1} match, these returns are not exactly comparable since the former is based on geometric compounding whereas the latter is not — as it is computed using an arithmetic average. This issue is mitigated by focusing on the *ex post* realised excess returns in (2.4) and (2.7), since they use the geometric and arithmetic average of the floating rate, respectively.

Figure 2.4: Unconditional *Ex Post* Excess Returns on US Overnight Indexed Swaps and Portfolios of Federal Funds Futures with Equivalent Horizon



Note: The left-hand column plots unconditional *ex post* excess returns for US OIS rates (red line) and portfolios of FFFs contracts (black line) calculated using equations (2.4) and (2.7), respectively. The right-hand column presents the differences between the two excess returns (blue line) at each date. The portfolios of FFFs contracts are constructed such that their horizon is equivalent to the horizon of corresponding-maturity OIS rates. The horizontal axis of each plot denotes the date of the *ex post* excess return and is labelled MM/YY. The data is plotted on penultimate business days of each month from January 2002 to December 2016.

for different monthly contract horizons $n = 1, 2, \dots, 11$.²⁰

All regressions are estimated using data observations on the penultimate business day of each calendar month. The sample runs from January 2002 to December 2016. The specific sample start date differs slightly across horizons due to the availability of US OIS rates at different tenors.²¹ The selection of maturities is determined by FFFs rate data availability.²²

Because contract horizons at adjacent time periods overlap, I compute standard errors using the heteroskedasticity and autocorrelation consistent procedure of Hodrick (1992), which generalises the Hansen and Hodrick (1980) method for overlapping contracts to the case of heteroskedasticity. Throughout the chapter, I report *t*-statistics based on these standard errors. In regressions (2.8) and (2.9), α_n^{OIS} and $\alpha_n^{FFF, port}$ denote the average *ex post* excess return on OIS contracts and portfolios of FFFs respectively. If these are insignificantly different from

²⁰Lloyd (2017b, Appendix B) also provides estimates of excess returns on *n*-month ahead (‘pure’) FFFs contracts using the regression:

$$rx_{t,t+n}^{FFF} = \alpha_n^{FFF} + \varepsilon_{t+n}^{FFF}$$

to check that the data transformation used to construct hypothetical portfolios of FFFs contracts does not influence the results.

²¹See appendix A.1 for a complete list of data sources and OIS rate availability.

²²Although available prior to 2002, detailed and regular series for 12 to 36-month ahead FFFs rates are not available for the whole post-2002 period.

zero, a contract is said to provide an accurate measure of expected future short-term interest rates. Because the OIS reference rate is a short-term money market rate, this empirical strategy formally analyses OIS rates as measures of short-term interest rate expectations, an indicator of the *de facto* monetary policy stance, rather than official policy rate settings.

Table 2.1 presents the baseline results from regressions (2.8) and (2.9) for the whole sample. Point estimates for average *ex post* realised excess returns are presented for 1 to 11-month OIS and portfolios of FFFs contracts using observations on the penultimate business day of each month.

The primary result is that 1 to 11-month OIS contracts provide measures of investors' interest rate expectations that are as good as those from corresponding-maturity portfolios of FFFs contracts, corroborating with figure 2.4. In fact, for the 3 to 11-month tenors, the point estimate of the average *ex post* excess return on US OIS contracts is marginally smaller than that on the corresponding-maturity portfolio of FFFs contracts. For example, the average *ex post* excess return on the 8-month OIS contract is 11.43 basis points, 0.76 basis points lower than the point estimate for the corresponding-maturity portfolio of FFFs contracts. Although the average *ex post* excess returns on 1 to 6-month OIS contracts are significant at the 10% level, at least, the same is true for 1 to 7-month portfolios of FFFs contracts. Moreover, in all cases, the average *ex post* excess returns are small — less than 9 basis points at the 6-month horizon — and subsequent analysis indicates that much of this can be explained by *ex ante* unexpected US monetary policy loosening in 2008.

A secondary result is that point estimates of average excess returns on OIS contracts, as well as portfolios of FFFs contracts, are increasing with the maturity of the contracts. This is consistent with the view that longer-horizon OIS rates may contain some term premia (Michaud and Upper, 2008).

Accounting for the 2008 US Monetary Policy Loosening During and after the 2007-2008 financial crisis, monetary policymakers lowered policy rates to their ELB. This was broadly unanticipated *ex ante* and, for this reason, may bias upwards the *ex post* excess returns presented in table 2.1. That is, the positive *ex post* excess returns in table 2.1 may actually reflect the unexpected policy loosening that occurred in response to the financial crisis and associated recession, rather than an excess return that reflects risk premia in OIS or FFFs contracts.

For this reason, I re-estimate (2.8) and (2.9) and include an additional dummy explanatory variable. This dummy is set to unity for periods in which the contracts measure expectations of future short-term interest rates during 2008, and zero otherwise. This aligns with the period in which US monetary policy was quickly loosened in response to the financial crisis and the associated recession.²³ For 1-month contracts, the dummy is set to unity from December 2007 to December 2008 (inclusive); for 2-month contracts, the dummy is set to unity from November 2007 to December 2008 (inclusive); and so on. Consequently, with the dummy variable included in the regression, α_n^{OIS} and $\alpha_n^{FFF,port}$ represent average *ex post* excess returns on OIS and portfolios of FFFs outside the 2008 period, respectively. The dummy variable coefficient captures the average additional increase in *ex post* excess returns attributable to the unpredicted nature of the 2007-2008 financial crisis and associated monetary policy loosening during 2008.

The results from these extended regressions are reported in table 2.2. The point estimates of

²³The federal funds rate target fell from 4.75% at the start of 2008 to 0-0.25% in December 2008.

Table 2.1: Average *Ex Post* Excess Returns on US OIS Contracts and Portfolios of FFFs Contracts of Comparable Maturity

Panel A: US OIS Contracts						
Maturity in Months	1	2	3	4	5	6
$\hat{\alpha}_n^{OIS}$	1.35***	2.52**	3.81**	5.22*	6.73*	8.45*
[<i>t</i> -statistic]	[2.72]	[2.19]	[1.98]	[1.92]	[1.86]	[1.81]
Maturity in Months	7^a	8^a	9	10^a	11^a	
$\hat{\alpha}_n^{OIS}$	9.89	11.43	14.03	15.05	17.16	
[<i>t</i> -statistic]	[1.64]	[1.52]	[1.56]	[1.39]	[1.36]	
Panel B: Portfolios of FFFs Contracts						
Maturity in Months	1	2	3	4	5	6
$\hat{\alpha}_n^{FFF,port}$	1.22***	2.44**	3.82*	5.30*	6.85*	8.52*
[<i>t</i> -statistic]	[2.39]	[2.01]	[1.91]	[1.88]	[1.83]	[1.76]
Maturity in Months	7	8	9	10	11	
$\hat{\alpha}_n^{FFF,port}$	10.36*	12.19	14.16	17.06	19.05	
[<i>t</i> -statistic]	[1.68]	[1.60]	[1.53]	[1.51]	[1.43]	

Note: Panel A reports results from regression (2.8) and panel B reports results from regression (2.9) for contracts/portfolios with 1-11 month maturity. Sample: Monthly Frequency, January 2002 to December 2016, but for those indicated by ^a May 2002 to December 2016 (due to OIS rate availability). Hodrick (1992) *t*-statistics are reported in square brackets. An excess return is significantly different from zero at the 1%, 5% and 10% significance level when the absolute value of the *t*-statistic exceeds 2.33, 1.96, 1.645 respectively. These are denoted with asterisks ***, ** and * for the 1%, 5% and 10% significance levels respectively. All figures are reported in basis points to two decimal places.

$\hat{\alpha}_n^{OIS}$ and $\hat{\alpha}_n^{FFF,port}$ are noticeably smaller than in table 2.1, typically at least half as small. The headline conclusion from table 2.1 is robust to the inclusion of a 2008 dummy. For 1 to 11-month OIS contracts, average *ex post* excess returns for the non-2008 period are quantitatively small, and the coefficients are statistically insignificant for the 5 to 11-month tenors. They are also quantitatively smaller than corresponding-maturity FFFs contracts. For example, the average *ex post* excess return for the 9-month OIS contract outside the 2008 period is 3.32 basis points, 10.71 basis points lower than its average *ex post* excess return for the whole 2002-2016 period and 0.94 basis points lower than the average *ex post* excess return on the comparable-maturity portfolio of FFFs contracts outside the 2008 period. Although the average *ex post* excess returns on OIS contracts for the non-2008 period are statistically significant at the 5% level for the 1 to 3-month tenors, and at the 10% level for the 4-month contract, their magnitude is small; the point estimate of the average *ex post* excess return on the 4-month OIS contract for the non-2008 period is just 1.54 basis points, 0.62 basis points lower than the comparable-maturity portfolio of FFFs. Thus, the usefulness of 1 to 11-month OIS rates as market-based measures of monetary policy expectations remains comparable to FFFs rates outside of the 2008 financial crisis period.

The coefficients on the 2008 dummy highlight the *ex ante* unpredicted nature of the financial crisis and associated policy loosening. For the OIS regressions in panel A of table 2.2, the coefficients on the 2008 dummy are positive and statistically significant, at the 5% level at least, for the 1 to 11-month OIS contracts. Moreover, the 2008 dummy coefficients are broadly increasing in the maturity of the OIS contracts, indicating that the extent to which the 2008 policy loosening was unanticipated increases at longer horizons. Interestingly, although the coefficients on the 2008 dummy for the FFFs regressions in panel B follow a similar qualitative pattern,

Table 2.2: *Ex Post* Excess Returns on US OIS Contracts and Portfolios of FFFs Contracts After Controlling for a 2008 Dummy

Panel A: US OIS Contracts						
Maturity in Months	1	2	3	4	5	6
$\hat{\alpha}_n^{OIS}$	0.50**	0.86**	1.14**	1.54*	1.91	2.22
[<i>t</i> -statistic]	[1.98]	[2.22]	[1.96]	[1.76]	[1.55]	[1.35]
2008 Dummy	12.76**	22.70**	33.70***	43.14***	52.72***	63.78***
[<i>t</i> -statistic]	[2.33]	[2.11]	[2.44]	[3.34]	[7.39]	[7.54]
Maturity in Months	7^a	8^a	9	10^a	11^a	
$\hat{\alpha}_n^{OIS}$	3.49	3.58	3.32	3.86	3.96	
[<i>t</i> -statistic]	[1.24]	[1.07]	[0.81]	[0.83]	[0.72]	
2008 Dummy	60.11***	69.37***	91.54***	88.47***	99.03***	
[<i>t</i> -statistic]	[6.21]	[7.13]	[10.26]	[9.39]	[6.73]	
Panel B: Portfolios of FFFs Contracts						
Maturity in Months	1	2	3	4	5	6
$\hat{\alpha}_n^{FFF,port}$	0.54*	1.18*	1.92*	2.16*	2.54	2.93
[<i>t</i> -statistic]	[1.82]	[1.79]	[1.70]	[1.64]	[1.43]	[1.32]
2008 Dummy	9.71*	16.56	22.92	35.20***	44.95***	54.56***
[<i>t</i> -statistic]	[1.77]	[1.44]	[1.53]	[2.43]	[4.82]	[6.25]
Maturity in Months	7	8	9	10	11	
$\hat{\alpha}_n^{FFF,port}$	3.49	3.98	4.26	7.43	9.84	
[<i>t</i> -statistic]	[1.24]	[1.14]	[1.00]	[1.11]	[1.07]	
2008 Dummy	63.01***	71.33***	81.16***	73.31***	65.26***	
[<i>t</i> -statistic]	[5.00]	[6.06]	[12.03]	[4.25]	[2.93]	

Note: Panel A reports results from regression (2.8) and panel B reports results from regression (2.9) with an additional dummy variable set equal to one on dates when a contract's horizon spans the year 2008. Sample: Monthly Frequency, January 2002 to December 2016, but for those indicated by ^a May 2002 to December 2016 (due to OIS rate availability). Hodrick (1992) *t*-statistics are reported in square brackets. An excess return is significantly different from zero at the 1%, 5% and 10% significance level when the absolute value of the *t*-statistic exceeds 2.33, 1.96, 1.645 respectively. These are denoted with asterisks ***, ** and * for the 1%, 5% and 10% significance levels respectively. All figures are reported in basis points to two decimal places.

the coefficients are statistically insignificant for 2 to 3-month portfolios of FFFs contracts. This implies that, at these maturities, FFFs rates did not include any statistically significant additional *ex post* excess returns during the 2008 period relative to the non-2008 period, indicating that the quality of these FFFs tenors as predictors of the future monetary policy stance did not significantly change during that period.

Overall, the results in tables 2.1 and 2.2 indicate that 1 to 11-month OIS rates provide measures of investors' interest rate expectations that are at least as good as those from comparable-horizon FFFs. This is important because FFFs have long been used to forecast future monetary policy, yet because OIS rates are available at longer maturities than FFFs and in a range of countries, OIS rates potentially offer a globally comparable financial market measure of monetary policy expectations.

2.4 OIS Rates from a Global Perspective

Here, I assess whether OIS rates accurately measure monetary policy expectations globally.

2.4.1 US OIS Contracts

I calculate unconditional *ex post* excess returns on US OIS contracts at a daily frequency, between January 2002 and December 2016, for the following tenors: 1 to 12 months; 15, 18 and 21 months; 2, 3, 4 and 5 years.²⁴ I estimate (2.8) using this daily frequency data. Again, *t*-statistics use heteroskedasticity and autocorrelation robust Hodrick (1992) standard errors.

Panel A of table 2.3 presents the results from these regressions. Daily frequency average *ex post* excess returns on 1 to 11-month OIS contracts are similar to the monthly frequency point estimates reported in table 2.1, indicating that monthly frequency estimates in section 2.3 are not susceptible to calendar effects. For example, the average *ex post* excess return on the 9-month OIS contract calculated with daily frequency data is 14.09 basis points, while the corresponding point estimate calculated using observations on the penultimate business day of each month is just 0.06 basis points lower at 14.03 basis points. The average *ex post* excess returns on the 1 to 5-month OIS contracts are significant at the 10% level for the whole 2002-2016 sample, but remain quantitatively small — not exceeding 6.87 basis points.

The 12 to 21-month OIS contracts, for which section 2.3 did not present monthly frequency results, exhibit statistically insignificant *ex post* excess returns for the 2002-2016 sample period. As with the 1 to 11-month tenors, the estimated average *ex post* excess returns on the 12 to 21-month contracts are increasing in the contract horizon, but remain insignificantly different from zero. At the 2-year horizon, the average *ex post* excess return is 53.27 basis points — over double the excess return on the 1-year contract of 20.98 basis points — and is statistically significant at the 10% level. In subsequent sensitivity analysis, I conclude that this marginal significance is primarily driven by the money market turmoil and unanticipated loosening of monetary policy during 2007-2008, as opposed to risk premia within the contract.

At longer horizons — 3, 4 and 5 years — OIS contracts have statistically significant positive *ex post* excess returns at the 1% level, indicative of term premia in longer-horizon OIS rates that blur their use as market-based measures of monetary policy expectations.

Accounting for 2008 US Monetary Policy Loosening As a sensitivity test, I regress the daily frequency unconditional *ex post* excess returns of OIS contracts on a constant and a 2008 dummy. As in section 2.3, this dummy is defined to capture the unanticipated nature of US monetary policy loosening in the wake of the 2007-2008 financial crisis. This regression accounts for the possibility that the unanticipated 2008 monetary policy loosening biases estimates of average *ex post* excess returns on OIS contracts upwards. That is, the positive average *ex post* excess returns presented in panel A of table 2.3 may actually reflect the *ex ante* unexpected nature of monetary policy accommodation during 2008, rather than risk premia in OIS contracts that complicate their use as a market-based measure of monetary policy expectations. As before, I define a 2008 dummy for each OIS contract maturity, set equal to unity on dates where the horizon of an OIS contract overlaps with the 2008 US policy accommodation. I define the policy accommodation period as January 22, 2008 — the first date on which the US policy rate was lowered in 2008 — to December 16, 2008 — the date on which the federal funds rate target was lowered to 0-0.25%.²⁵ Here, the estimated α_n^{OIS} coefficient can be interpreted as the average *ex*

²⁴The choice of tenors is determined by data availability. See appendix A.1.

²⁵For example, the dummy is set to 1 on days between December 22, 2007 and December 16, 2008 (inclusive) and zero otherwise for the 1-month contract.

Table 2.3: Average *Ex Post* Excess Returns on US OIS Contracts at Daily Frequency

Panel A: US OIS Contracts					
Maturity in Months	1	2	3	4	5
$\hat{\alpha}_n^{OIS}$	1.34*	2.54*	3.86*	5.32*	6.87*
[<i>t</i> -statistic]	[1.79]	[1.79]	[1.81]	[1.79]	[1.69]
Maturity in Months	6	7^a	8^a	9	10^a
$\hat{\alpha}_n^{OIS}$	8.54	9.93	11.54	14.09	15.20
[<i>t</i> -statistic]	[1.56]	[1.39]	[1.30]	[1.34]	[1.22]
Maturity in Months	11^a	12	15	18	21
$\hat{\alpha}_n^{OIS}$	17.24	20.98	27.61	35.59	44.25
[<i>t</i> -statistic]	[1.20]	[1.32]	[1.33]	[1.43]	[1.60]
Maturity in Months	24	36^b	48^b	60^b	
$\hat{\alpha}_n^{OIS}$	53.27*	86.86***	126.32***	173.60***	
[<i>t</i> -statistic]	[1.86]	[4.61]	[5.37]	[8.48]	
Panel B: US OIS Contracts with 2008 Dummy					
Maturity in Months	1	2	3	4	5
$\hat{\alpha}_n^{OIS}$	0.41*	0.76*	1.14*	1.50	1.90
[<i>t</i> -statistic]	[1.75]	[1.80]	[1.75]	[1.61]	[1.40]
2008 Dummy	14.11	24.65**	34.79***	45.33***	54.76***
[<i>t</i> -statistic]	[1.58]	[2.00]	[3.08]	[10.78]	[5.04]
Maturity in Months	6	7^a	8^a	9^a	10^a
$\hat{\alpha}_n^{OIS}$	2.22	1.88	1.86	3.28	2.41
[<i>t</i> -statistic]	[1.19]	[0.83]	[0.60]	[0.72]	[0.48]
2008 Dummy	65.20***	76.26***	86.01***	93.12***	101.68***
[<i>t</i> -statistic]	[8.34]	[14.75]	[10.53]	[9.79]	[9.78]
Maturity in Months	11^a	12	15	18	21
$\hat{\alpha}_n^{OIS}$	2.93	5.73	8.34	12.11	16.02
[<i>t</i> -statistic]	[0.48]	[0.68]	[0.65]	[0.74]	[0.84]
2008 Dummy	108.41***	112.22***	123.26***	131.95***	141.13***
[<i>t</i> -statistic]	[10.56]	[13.46]	[11.47]	[7.16]	[6.00]
Maturity in Months	24	36^b	48^b	60^b	
$\hat{\alpha}_n^{OIS}$	20.91	43.47**	84.49***	155.05***	
[<i>t</i> -statistic]	[1.07]	[2.01]	[3.57]	[21.59]	
2008 Dummy	145.15***	132.62***	93.14***	31.15	
[<i>t</i> -statistic]	[4.77]	[55.52]	[3.07]	[0.99]	

Note: Results from regression (2.8) for daily frequency OIS contracts. Sample: Daily Frequency, January 1, 2002 to December 31, 2016, but for those indicated by ^a May 7, 2002 to December 31, 2016 and ^b February 14, 2002 to December 31, 2016 (due to OIS rate availability). Hodrick (1992) *t*-statistics are reported in square brackets. An excess return is significantly different from zero at the 1%, 5% and 10% significance level when the absolute value of the *t*-statistic exceeds 2.33, 1.96, 1.645 respectively. These are denoted with asterisks ***, ** and * for the 1%, 5% and 10% significance levels respectively. All figures are reported in basis points to two decimal places. The 2008 dummy is set equal to unity on dates where the OIS contract horizon overlaps with the January 22, 2008 to December 16, 2008 US monetary policy loosening.

post excess return on an n -month OIS contract in periods for which the contract's horizon did not overlap with the 2008 US monetary policy loosening. The coefficient on the 2008 dummy captures the additional increase in *ex post* excess returns during the 2008 period.

The results in panel B of table 2.3 further support the main conclusions of the chapter. In particular, the average *ex post* excess returns outside the 2008 period on 4 to 24-month OIS contracts are insignificantly different from zero, and are all substantially smaller than the point estimates of the average *ex post* excess return for the whole 2002-2016 sample. For example, the average *ex post* excess return on the 8-month contract for the non-2008 period is 1.86 basis points, approximately 1/6th of the corresponding point estimate for the whole 2002-2016 period. Although the average *ex post* excess returns on 1 to 3-month OIS contracts are significant at the 10% level, they are small — less than 1.14 basis points — indicating that OIS rates at these tenors do provide accurate measures of interest rate expectations.

The 2-year OIS contract exhibits an average *ex post* excess return of 20.91 basis points for the non-2008 period, a figure which is insignificantly different from zero. Thus, the (marginally) significant average figure for the whole 2002-2016 period appears to be a result of the unanticipated events of 2008, rather than risk premia in the contract.

At longer horizons — 3, 4 and 5 years — OIS contracts still exhibit significant term premia in the non-2008 period. However, the average *ex post* excess returns on these contracts for the non-2008 period are smaller than for the 2002-2016 period as a whole. For example, the non-2008 average *ex post* excess return for the 3-year contract is 43.47 basis points, which is approximately half the estimate for the whole 2002-2016 sample.

For the 2 to 48-month tenors, the 2008 dummy coefficients are all positive and significant, consistent with the claim that the 2008 monetary policy loosening was *ex ante* unexpected.

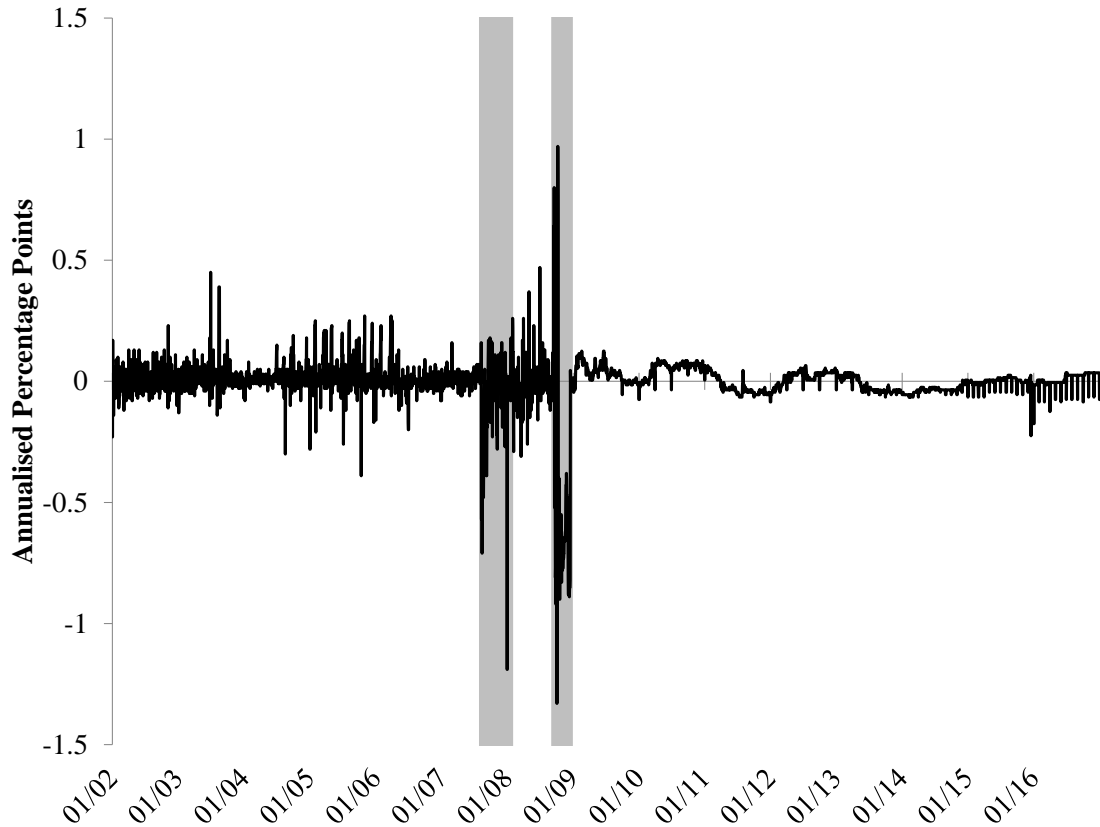
Accounting for 2007-2008 Money Market Turmoil Although the average *ex post* excess returns on 1 to 3-month OIS contracts in panel B of table 2.3 are small, they are significant at the 10% level. I conduct further sensitivity analysis by looking into the influence of 2007-2008 money market turmoil on estimated *ex post* excess returns. The money market turmoil of 2007-2008 had notable implications for US overnight interest rates. Taylor and Williams (2009) document that the effective federal funds rate, the OIS reference rate, jumped to unusually high levels compared with the Fed's target for the federal funds rate, the policy rate, on August 9, 2007.²⁶ On the following day, the Federal Reserve Bank of New York pumped liquidity into the market, leading to a marked fall in the effective federal funds rate relative to the federal funds target rate, as shown in figure 2.5.²⁷

In general, the impact of differences between reference and policy rates will be expected to diminish at longer horizons, evidenced by the insignificant *ex post* excess returns on 4 to 24-month contracts reported in panel B of table 2.3. However, figure 2.5 illustrates that large differences between the effective federal funds rate and the federal funds target rate occurred during the 2007-2008 period, following money market turbulence that erupted on August 9, 2007 and, after some recovery in money market conditions during the first half of 2008, reignited after the failure of Lehman Brothers in September 2008. The turbulence led to a dramatic change in money market conditions.

²⁶On August 9, 2007, the difference between the effective federal funds rate and its target rose from 2 to 16 basis points.

²⁷On August 10, 2007 the effective federal funds rate was 57 basis points below the federal funds target rate.

Figure 2.5: Effective Federal Funds Rate Minus the Federal Funds Target Rate



Note: This figure depicts the difference between the effective federal funds rate and the federal funds target rate at a daily frequency from January 2002 to December 2016. From December 16, 2008, the difference is calculated by assuming that the federal funds target rate was halfway between its lower and upper bounds. The grey areas denote the periods of money market turbulence. The first begins on August 9, 2007 and ends on January 22, 2008, when the federal funds target rate was cut by 75 basis points. The second begins on September 15, 2008 — the day of Lehman Brothers’ failure — and ends on December 16, 2008. *Data Source:* Federal Reserve.

The relationship between the effective federal funds rate and the federal funds rate target changed significantly during the period of initial money market turmoil, from August 9, 2007 to January 21, 2008 — the day before the Federal Open Market Committee cut the federal funds target rate by 75 basis points.²⁸ The effective federal funds rate was also significantly below its target following the failure of Lehman Brothers on September 15, 2008 until December 16, 2008 when the Fed cut its short-term policy rate to its ELB.²⁹

Because the money market turmoil had a significant influence on the short-term interest rates of direct relevance to OIS rates, I carry out further sensitivity analysis to account for the additional effect that this may have had on the *ex post* excess returns on OIS contracts. To do this, I estimate (2.8) with three dummy variables: (i) a dummy variable for the initial money market turmoil from August 9, 2007 to January 21, 2008; (ii) a dummy accounting for the unanticipated monetary policy loosening between January 22, 2008 and September 14, 2008;

²⁸The average difference between the effective federal funds rate and the federal funds target rate prior to the money market turmoil was 0 percent points. This figure is calculated using daily data from January 2, 2007 to August 8, 2007 ($N = 219$ observations), with a standard deviation of 0.03 percent points. The corresponding figure for the period of initial money market turmoil was -0.09 percent points. This figure is calculated using daily data from August 9, 2007 to January 21, 2008 ($N = 166$ observations), with standard deviation of 0.21 percent points. The corresponding t -statistic from a difference-in-mean hypothesis test is statistically significant.

²⁹The average difference between the effective federal funds rate and the federal funds target rate from September 15, 2008 to December 15, 2008 was -0.59 percent points.

Table 2.4: Average *Ex Post* Excess Returns on US OIS Contracts at Daily Frequency with Controls for 2008 Monetary Policy Loosening and 2007-2008 Money Market Turmoil

Maturity in Months	1	2	3	4	5
$\hat{\alpha}_n^{OIS}$	0.23	0.33	0.46	0.63	0.94
[<i>t</i> -statistic]	[0.88]	[0.85]	[0.87]	[0.77]	[0.75]
Initial Mon. Market Dummy	7.90	17.17***	24.18***	29.90***	34.04**
[<i>t</i> -statistic]	[1.59]	[3.04]	[3.07]	[2.39]	[2.17]
Mon. Pol. Loosening Dummy	0.18	3.56	7.86	13.27**	18.56***
[<i>t</i> -statistic]	[0.05]	[0.57]	[0.92]	[1.97]	[2.91]
Post-Lehman Dummy	35.95***	45.16***	51.46***	54.49***	55.02***
[<i>t</i> -statistic]	[3.49]	[6.43]	[4.42]	[7.93]	[5.23]
Maturity in Months	6	7^a	8^a	9	10^a
$\hat{\alpha}_n^{OIS}$	1.42	1.28	1.67	3.34	2.58
[<i>t</i> -statistic]	[0.76]	[0.56]	[0.52]	[0.71]	[0.50]
Initial Mon. Market Dummy	34.10**	31.67*	25.77	18.41	14.85
[<i>t</i> -statistic]	[2.11]	[1.84]	[1.52]	[1.30]	[1.28]
Mon. Pol. Loosening Dummy	29.24**	42.40**	56.31***	65.17***	71.62***
[<i>t</i> -statistic]	[2.10]	[2.31]	[2.77]	[3.37]	[3.95]
Post-Lehman Dummy	49.66***	44.17***	38.10***	39.91***	45.16***
[<i>t</i> -statistic]	[3.97]	[3.35]	[2.71]	[3.54]	[14.65]
Maturity in Months	11^a	12	15	18	21
$\hat{\alpha}_n^{OIS}$	3.10	5.96	7.91	11.64	16.51
[<i>t</i> -statistic]	[0.49]	[0.70]	[0.61]	[0.72]	[0.95]
Initial Mon. Market Dummy	12.59	8.29	8.95	-0.74	-24.42
[<i>t</i> -statistic]	[1.33]	[1.37]	[0.91]	[-0.05]	[-1.08]
Mon. Pol. Loosening Dummy	73.81***	75.83***	56.85***	48.59***	63.23***
[<i>t</i> -statistic]	[5.56]	[16.69]	[10.23]	[13.00]	[6.35]
Post-Lehman Dummy	52.97***	58.62***	98.33***	125.43***	134.73***
[<i>t</i> -statistic]	[4.17]	[3.97]	[8.67]	[6.63]	[9.28]
Maturity in Months	24	36^b	48^b	60^b	
$\hat{\alpha}_n^{OIS}$	22.31	47.23***	90.77***	165.10***	
[<i>t</i> -statistic]	[1.44]	[3.10]	[4.95]	[25.57]	
Initial Mon. Market Dummy	-45.19*	-108.98***	-129.48***	-144.30***	
[<i>t</i> -statistic]	[-1.70]	[-6.24]	[-7.75]	[-8.15]	
Mon. Pol. Loosening Dummy	70.49***	27.00	-4.38	-33.59***	
[<i>t</i> -statistic]	[8.18]	[1.34]	[-0.52]	[-3.89]	
Post-Lehman Dummy	145.22***	230.61***	231.30***	201.68***	
[<i>t</i> -statistic]	[11.17]	[11.76]	[14.18]	[32.61]	

Note: Results from regression (2.8) for daily frequency US OIS contracts. Sample: Daily Frequency, January 1, 2002 to December 31, 2016, but for those indicated by ^a May 7, 2002 to December 31, 2016 and ^b February 14, 2002 to December 31, 2016 (due to OIS rate availability). Hodrick (1992) *t*-statistics are reported in square brackets. An excess return is significantly different from zero at the 1%, 5% and 10% significance level when the absolute value of the *t*-statistic exceeds 2.33, 1.96, 1.645 respectively. These are denoted with asterisks ***, ** and * for the 1%, 5% and 10% significance levels respectively. All figures are reported in basis points to two decimal places. The 'Initial Mon. Market Dummy' is set equal to unity on dates where the OIS contract horizon overlaps with the August 9, 2007 to January 21, 2008 money market turmoil. The 'Mon. Pol. Loosening Dummy' is set equal to unity on dates where the OIS contract horizon overlaps with the January 22, 2008 to September 14, 2008 US monetary policy loosening. The 'Post-Lehman Dummy' is set equal to unity on dates where the OIS contract horizon overlaps with the September 15, 2008 to December 16, 2008 money market turmoil and monetary policy loosening following the failure of Lehman Brothers.

and (iii) a dummy to account for the money market turmoil and monetary policy loosening that occurred between September 15, 2008 and December 16, 2008. As for the 2008 dummy in panel B of table 2.3, the dummy variables are set to unity on dates when the horizon of an OIS contract overlaps with the stated period — not only on the day the *ex post* excess return is recorded. The results of these regressions are reported in table 2.4. The estimated α_n^{OIS} coefficient can be interpreted as the average *ex post* excess return on an n -month OIS contract in periods for which the contract’s horizon did not overlap with the 2007-2008 US money market turbulence and monetary policy loosening.

The main conclusions of the chapter are strengthened by the inclusion of the money market turmoil dummy. The average *ex post* excess returns on 1 to 24-month US OIS contracts are insignificantly different from zero outside the 2007-2008 period. At longer horizons — 3, 4 and 5 years — OIS contracts still exhibit statistically significant *ex post* excess returns. Moreover, the dummy variable coefficients indicate that the post-Lehman money market turmoil and monetary policy loosening typically had the strongest upward influence on OIS *ex post* excess returns.

Overall, the above results support two of the main conclusions in this chapter. First, 1 to 11-month OIS contracts provide measures of investors’ interest rate expectations as good as comparable-horizon FFFs contracts. Second, 1 to 24-month OIS contracts accurately predict the future path of monetary policy on average; at longer horizons, OIS rates include statistically significant *ex post* excess returns that reflect premia for risks in OIS contracts.

2.4.2 UK OIS Contracts

To assess the global usefulness of OIS rates as financial market-based measures of monetary policy expectations, I apply the empirical specification to UK OIS contracts. I calculate unconditional *ex post* excess returns on UK OIS contracts, from January 2001 to December 2016, at a daily frequency. I use UK OIS rates of the following maturities: 1, 2, 3, 4, 5, 6, 7, 8, 9, 10, 11, 12, 18 and 24 months.³⁰

To calculate unconditional *ex post* excess returns on UK OIS contracts, I make two alterations to the equations laid out in section 2.2. First, because calculations of OIS floating leg payments occur with no spot lag ($s = 0$), I calculate the floating interest rate from the trade date to maturity. Second, because UK OIS rates are quoted according to the *Actual 365* convention, (2.3) is a multiple of $365/N$, not $360/N$. With daily frequency *ex post* excess returns on UK OIS contracts, I estimate average *ex post* excess returns using (2.8).

Table 2.5 demonstrates that 2 to 18-month UK OIS contracts exhibit statistically insignificant average *ex post* excess returns for the whole sample period. Although the average *ex post* excess return on the 1-month contract is significant at the 1% level, it is small in magnitude at -2.72 basis points. Thus, 1 to 18-month UK OIS rates appear to provide accurate measures of interest rate expectations on average. As in the US market, the average *ex post* excess return on the 2-year UK OIS contract, of 47.15 basis points, is statistically significant (at the 10% level) for the whole 2001-2016 period.

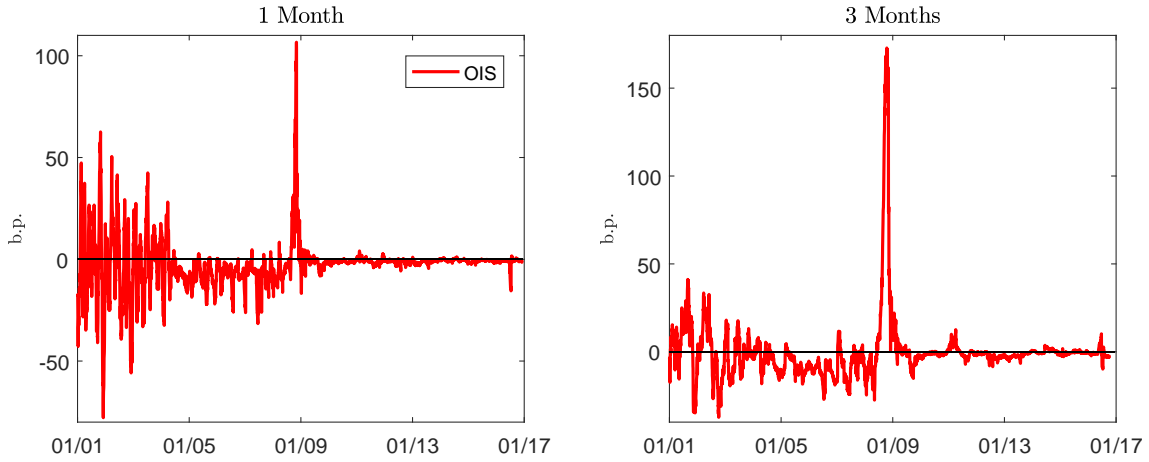
Accounting for UK Monetary Policy Loosening Figure 2.6 plots the unconditional *ex post* excess returns on 1 and 3-month UK OIS contracts. There is a notable spike in both in

³⁰The selection of maturities and sample length is, again, determined by data availability. See appendix A.1.

Table 2.5: Average *Ex Post* Excess Returns on UK OIS Contracts at Daily Frequency

Maturity in Months	1	2	3	4	5
$\hat{\alpha}_n^{OIS}$	-2.72***	-1.61	-0.32	1.14	2.73
[<i>t</i> -statistic]	[-3.33]	[-0.99]	[-0.12]	[0.31]	[0.59]
Maturity in Months	6	7	8	9	10
$\hat{\alpha}_n^{OIS}$	4.46	6.44	8.46	10.61	12.91
[<i>t</i> -statistic]	[0.78]	[0.93]	[1.02]	[1.09]	[1.16]
Maturity in Months	11	12	18	24	
$\hat{\alpha}_n^{OIS}$	15.37	17.77	30.72	47.15*	
[<i>t</i> -statistic]	[1.22]	[1.27]	[1.48]	[1.77]	

Note: Results from regression (2.8) for UK OIS contracts. Sample: January 1, 2001 to December 31, 2016, Daily Frequency. Hodrick (1992) *t*-statistics are reported in square brackets. An excess return is significantly different from zero at the 1%, 5% and 10% significance level when the absolute value of the *t*-statistic exceeds 2.33, 1.96, 1.645 respectively. These are denoted with asterisks ***, ** and * for the 1%, 5% and 10% significance levels respectively. All figures are reported in basis points to two decimal places.

Figure 2.6: Unconditional *Ex Post* Excess Returns on UK Overnight Indexed Swaps

Note: Time series of unconditional *ex post* excess returns for UK OIS rates calculated using equation (2.4). Sample: January 1, 2001 to December 31, 2016; daily frequency. The horizontal axis of each plot denotes the date of the *ex post* excess return and is labelled MM/YY.

late-2008, when UK monetary policy was being loosened in response to the financial crisis.³¹

I investigate the significant *ex post* excess returns on 1-month and 2-year OIS contracts in table 2.5 by conducting sensitivity analysis to account for the possibility that the *ex ante* unanticipated UK monetary policy accommodation in response to the 2007-2008 financial crisis biases estimates of average *ex post* excess returns. I regress the unconditional *ex post* excess returns on UK OIS contracts on a constant and a monetary policy loosening dummy variable, set equal to unity on dates where the OIS contract horizon overlaps with the period spanning December 6, 2007 to March 5, 2009. These dates are chosen, as following the onset of financial market turmoil in the summer 2007, the Bank of England cut Bank Rate from 5.75% to 0.5% between December 6, 2007 to March 5, 2009. I do not include a dummy variable to separately account for the potential effects of money market turmoil for two reasons. First, the relationship between SONIA and Bank Rate was not significantly different during the money market turmoil

³¹The *ex post* excess returns in figure 2.6 are also highly volatile from 2001 to mid-2004, a consequence of UK money market operating procedures at the time. I study this in subsequent robustness analysis.

Table 2.6: Average *Ex Post* Excess Returns on UK OIS Contracts at Daily Frequency with Controls for UK Monetary Policy Loosening

Maturity in Months	1	2	3	4	5
$\hat{\alpha}_n^{OIS}$	-3.24***	-2.94***	-2.55***	-2.15	-1.71
[<i>t</i> -statistic]	[-5.30]	[-3.61]	[-2.54]	[-1.64]	[-0.93]
Mon. Pol. Loosening Dummy	6.28	14.78	23.39	32.27	41.22
[<i>t</i> -statistic]	[0.95]	[1.06]	[1.15]	[1.34]	[1.62]
Maturity in Months	6	7	8	9	10
$\hat{\alpha}_n^{OIS}$	-1.12	-0.34	0.51	1.52	2.65
[<i>t</i> -statistic]	[-0.44]	[-0.10]	[0.12]	[0.30]	[0.45]
Mon. Pol. Loosening Dummy	49.25*	56.51**	63.03***	68.47***	74.04***
[<i>t</i> -statistic]	[1.92]	[2.20]	[2.38]	[2.45]	[2.65]
Maturity in Months	11	12	18	24	
$\hat{\alpha}_n^{OIS}$	3.91	5.35	13.15	25.26*	
[<i>t</i> -statistic]	[0.57]	[0.71]	[1.15]	[1.71]	
Mon. Pol. Loosening Dummy	79.27***	82.62***	92.30***	94.59***	
[<i>t</i> -statistic]	[3.02]	[3.31]	[3.71]	[2.37]	

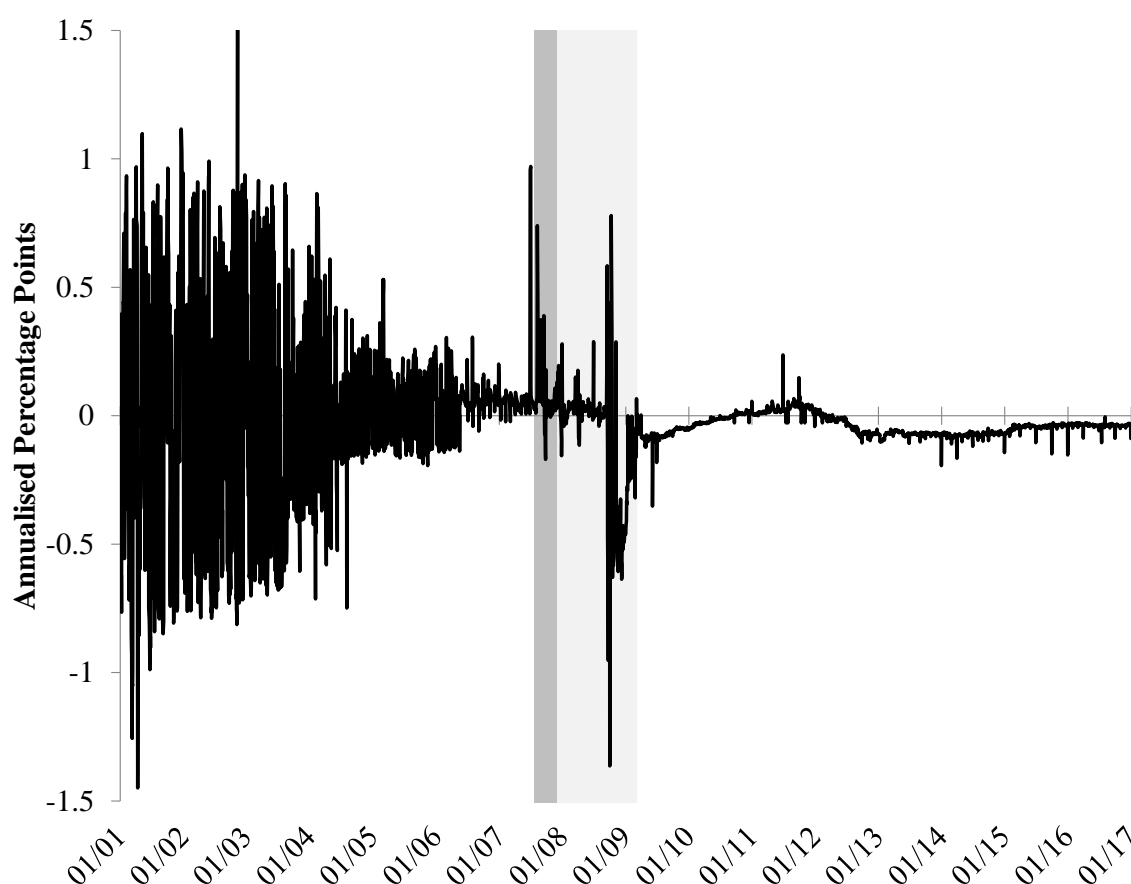
Note: Results from regression (2.8) for daily frequency UK OIS contracts. Hodrick (1992) *t*-statistics are reported in square brackets. An excess return is significantly different from zero at the 1%, 5% and 10% significance level when the absolute value of the *t*-statistic exceeds 2.33, 1.96, 1.645 respectively. These are denoted with asterisks ***, ** and * for the 1%, 5% and 10% significance levels respectively. All figures are reported in basis points to two decimal places. Sample: January 1, 2001 to December 31, 2016; daily frequency; the ‘Mon. Pol. Loosening Dummy’ is set equal to unity on dates where the OIS contract horizon overlaps with the December 6, 2007 to March 5, 2009 UK monetary policy loosening.

relative to the months prior to it.³² Second, figure 2.6 shows that there is no notable spike in unconditional *ex post* excess returns on UK OIS contracts around initial money market turmoil in summer 2007, in contrast to the US excess returns in figure 2.4, indicating that UK money market conditions did not appreciably influence the variable of interest at this time.

Table 2.6 presents the regression results. The point estimates of α_n^{OIS} can be interpreted as the average *ex post* excess return on an *n*-month OIS contract in periods for which the contract’s horizon did not overlap with UK monetary policy loosening. In comparison to table 2.5, table 2.6 indicates that fewer OIS tenors have statistically insignificant *ex post* excess returns. The estimated α_n^{OIS} coefficients for the 1 to 3-month contracts are negative and statistically significant, like the 1-month contract coefficient in table 2.5. Although the point estimates are quantitatively small — less than 3.25 basis points in absolute value — they indicate that these OIS tenors may not have provided the most accurate measures of UK interest rate expectations outside the 2007-2009 period. Moreover, while the estimated α_n^{OIS} coefficient for the 24-month contract is almost half its value in table 2.5, it remains significant at the 10% level.

³²The average difference between SONIA and the UK Bank Rate prior to the money market turmoil was 0.053 percent points. This figure is calculated using daily data from January 2, 2007 to August 8, 2007, excluding a one-day spike on June 29 ($N = 152$ observations), with a standard deviation on 0.039 percent points. The corresponding figure for the initial period of money market turmoil was 0.075 percent points. This figure is calculated using daily data from August 9, 2007 to December 5, 2007 ($N = 84$ observations), with standard deviation of 0.126 percent points. The corresponding *t*-statistic from a difference-in-mean hypothesis test is -1.6, implying the difference is statistically insignificant. Nevertheless, when the regression is estimated with an additional money market turmoil dummy (set equal to unity when contracts mature between August 9, 2007 and December 5, 2007), $\hat{\alpha}_n^{OIS}$ coefficients remain significant at the 1% level for 1 to 3-month contracts, while the 24-month coefficient is significant at 1% level. All other $\hat{\alpha}_n^{OIS}$ coefficients remain statistically insignificant.

Figure 2.7: SONIA Minus UK Bank Rate



Note: This figure depicts the difference between the SONIA and the UK Bank Rate at a daily frequency from January 2001 to December 2016. The left-hand dark grey area denotes the first period of money market turbulence in 2007 — August 9 to December 5, 2007. The right-hand light grey area denotes the period of monetary policy loosening — December 6, 2007 to March 5, 2009. *Data Source:* Bank of England.

Accounting for UK Money Market Reform One potential explanation for the significant *ex post* excess returns at the 1 to 3 and 24-month horizons in table 2.6 relates to differences between the OIS reference rate and the headline policy rate (Shareef, 2013). In the UK, open market operations are used to try and minimise the difference between Bank Rate, the policy rate, and SONIA, the OIS reference rate. Figure 2.7 plots this difference. It illustrates that, from 2001 to mid-2004, differences were extremely large, peaking at almost 1.5 percentage points in late-2002. The largest absolute difference between the effective federal funds rate and the federal funds target rate in the 2002-2016 period was less than a third of this (figure 2.5). Bank of England (2004) acknowledge that, during this period, sterling overnight rates were considerably more volatile than for other countries. Figure 2.6 demonstrates that *ex post* excess returns on OIS contracts were also highly volatile at this time, mirroring movements in figure 2.7. The differences between SONIA and Bank Rate were narrowed by two changes to the Bank of England's money market operations on July 22, 2004,³³ and May 17, 2006.³⁴

To assess the extent to which volatility of overnight rates between 2001 and July 2004 can explain the significant *ex post* excess returns on 1 to 3 and 24-month OIS contracts, I regress the

³³See: www.bankofengland.co.uk/archive/Documents/historicpubs/news/2004/082.pdf.

³⁴See: www.bankofengland.co.uk/archive/Documents/historicpubs/news/2006/055.pdf.

unconditional *ex post* excess returns on UK OIS contracts on a constant, the monetary policy loosening dummy used in table 2.6, and a money market volatility dummy. The money market volatility dummy is set equal to unity on dates where the OIS contract overlaps with the period spanning January 1, 2001 to July 22, 2004, the date of initial UK money market reform. This reform substantially reduced the volatility of UK overnight rates.

The results are presented in panel A of table 2.7. Money market volatility from 2001 to 2004 does help to explain the significance of the 24-month tenor in tables 2.5 and 2.6. In panel A of table 2.7, the average *ex post* excess return on the 24-month contract is statistically insignificant at 21.47 basis points. However, the money market volatility dummy does not reverse the significance of the short-horizon excess returns. Although quantitatively small, the average *ex post* excess returns on the 1 to 5-month contracts, outside of the periods covered by the two dummy variables, are significantly negative.

Panel B of table 2.7 assesses whether the significantly negative *ex post* excess returns on short-horizon contracts have changed over time. Here, I estimate the unconditional average *ex post* excess return on 1 to 5-month contracts for the post-2008 period only, beginning the sample on April 1, 2009. Although the 1 to 3-month $\hat{\alpha}_n^{OIS}$ estimates are significant at the 1% level, they are quantitatively small. Average *ex post* excess returns on 1 to 5-month contracts are less than 1.20 basis points in magnitude, around a third of their size in panel A of table 2.7. The fact these excess returns are smaller in magnitude in this period is consistent with a reduction in UK money market volatility and a narrowing of the spread between SONIA and Bank Rate following changes to sterling money market operating procedures in March 2009 (Jackson and Sim, 2013; Osborne, 2016). Thus, the results indicate that the accuracy of very short-horizon contracts as measures of UK interest rate expectations have improved since the financial crisis. This is further supported in figure 2.6, where excess returns are visibly close to, but slightly below, zero for most of the period after the financial crisis, following sizeable spikes in late-2008.

2.4.3 Eurozone OIS Contracts

EURONIA is the overnight floating reference rate used to calculate *ex post* excess returns on Eurozone OIS contracts. As per market convention, the contracts have a two-day spot lag and obey the *Actual 360* dating norm. I use Eurozone OIS rates, between January 2000 and December 2016, of the following maturities: 1, 2, 3, 4, 5, 6, 7, 8, 9, 10, 11, 12, 15, 18, 24 and 36 months.³⁵

Table 2.8 presents the estimated average *ex post* excess returns on Eurozone OIS contracts for the whole 2000-2016 sample. These unconditional average *ex post* returns are significant at all horizons, at the 10% level at least. Although the average *ex post* excess returns at short-term horizons are small — the 1-month point estimate is just 1.12 basis points — there is no distinction in the significance of short and long-term horizon tenors. At first sight, these findings challenge the claim that 1 to 24-month Eurozone OIS rates provide useful measures of investors' expectations of future short-term interest rates. However, this average result for the whole 2000-2016 sample masks variation within the period.

Accounting for Eurozone Money Market Developments from August 2007 To investigate these results, I conduct sensitivity analysis to assess whether developments in Eurozone

³⁵The selection of maturities is, again, determined by data availability — see appendix A.1.

Table 2.7: Average *Ex Post* Excess Returns on UK OIS Contracts at Daily Frequency with Controls for UK Monetary Policy Loosening and UK Money Market Volatility

Panel A: UK OIS Contracts with Two Dummy Variables					
Maturity in Months	1	2	3	4	5
$\hat{\alpha}_n^{OIS}$	−3.29***	−3.44***	−3.40***	−3.34***	−3.30*
[<i>t</i> -statistic]	[−6.00]	[−4.39]	[−3.44]	[−2.58]	[−1.93]
Mon. Mkt. Volatility Dummy	0.17	2.05	3.42	4.74	6.31
[<i>t</i> -statistic]	[0.09]	[0.88]	[1.30]	[1.49]	[1.36]
Mon. Pol. Loosening Dummy	6.32	15.28	24.23	33.45	42.81*
[<i>t</i> -statistic]	[0.96]	[1.10]	[1.19]	[1.39]	[1.68]
Maturity in Months	6	7	8	9	10
$\hat{\alpha}_n^{OIS}$	−3.13	−2.80	−2.47	−1.96	−1.40
[<i>t</i> -statistic]	[−1.39]	[−0.97]	[−0.70]	[−0.47]	[−0.28]
Mon. Mkt. Volatility Dummy	7.81	9.51	11.37	13.16	15.12
[<i>t</i> -statistic]	[1.13]	[1.02]	[0.96]	[0.94]	[0.94]
Mon. Pol. Loosening Dummy	51.25**	58.96**	66.00***	71.96***	78.09***
[<i>t</i> -statistic]	[2.01]	[2.32]	[2.54]	[2.64]	[2.90]
Maturity in Months	11	12	18	24	
$\hat{\alpha}_n^{OIS}$	−0.67	0.41	7.71	21.47	
[<i>t</i> -statistic]	[−0.12]	[0.06]	[0.64]	[1.42]	
Mon. Mkt. Volatility Dummy	16.90	17.91	18.17	11.62	
[<i>t</i> -statistic]	[0.94]	[0.91]	[1.03]	[−0.69]	
Mon. Pol. Loosening Dummy	83.85***	87.57***	97.75***	98.39***	
[<i>t</i> -statistic]	[3.41]	[3.89]	[3.10]	[2.27]	
Panel B: UK Short-Horizon OIS Contracts with Post-2008 Sample					
Maturity in Months	1	2	3	4	5
$\hat{\alpha}_n^{OIS}$	−1.08***	−1.20***	−1.14***	−0.91	−0.52
[<i>t</i> -statistic]	[−7.03]	[−3.87]	[−2.33]	[−1.39]	[−0.62]

Note: Results from regression (2.8) for daily frequency UK OIS contracts. Hodrick (1992) *t*-statistics are reported in square brackets. An excess return is significantly different from zero at the 1%, 5% and 10% significance level when the absolute value of the *t*-statistic exceeds 2.33, 1.96, 1.645 respectively. These are denoted with asterisks ***, ** and * for the 1%, 5% and 10% significance levels respectively. All figures are reported in basis points to two decimal places. **Panel A:** January 1, 2001 to December 31, 2016 sample; daily frequency; the ‘Mon. Pol. Loosening Dummy’ is set equal to unity on dates where the OIS contract horizon overlaps with the December 6, 2007 to March 5, 2009; ‘Mon. Mkt. Volatility Dummy’ is set equal to unity on dates where OIS contract horizon overlaps with January 1, 2001 to July 22, 2004. **Panel B:** April 1, 2009 to December 31, 2016 sample; daily frequency.

Table 2.8: Average *Ex Post* Excess Returns on Eurozone OIS Contracts at Daily Frequency

Maturity in Months	1	2	3	4	5
$\hat{\alpha}_n^{OIS}$	1.12***	2.24***	3.51**	4.98**	6.61**
[<i>t</i> -statistic]	[2.48]	[2.55]	[2.17]	[2.06]	[2.02]
Maturity in Months	6	7	8	9	10
$\hat{\alpha}_n^{OIS}$	8.44**	10.36**	12.41**	14.57**	16.84**
[<i>t</i> -statistic]	[2.01]	[2.00]	[2.00]	[1.99]	[1.98]
Maturity in Months	11	12	15^a	18	21^a
$\hat{\alpha}_n^{OIS}$	19.19**	21.64*	28.17**	35.29**	43.15***
[<i>t</i> -statistic]	[1.98]	[1.99]	[2.05]	[2.23]	[2.48]
Maturity in Months	24	36^b			
$\hat{\alpha}_n^{OIS}$	51.26***	80.08***			
[<i>t</i> -statistic]	[2.83]	[5.00]			

Note: Results from regression (2.8) for Eurozone OIS contracts. Sample: Daily Frequency, January 3, 2000 to December 31, 2016, but for those indicated by ^a August 22, 2001 to December 31, 2016 and ^b March 3, 2004 to December 31, 2016 (due to OIS rate availability). Hodrick (1992) *t*-statistics are reported in square brackets. An excess return is significantly different from zero at the 1%, 5% and 10% significance level when the absolute value of the *t*-statistic exceeds 2.33, 1.96, 1.645 respectively. These are denoted with asterisks ***, ** and * for the 1%, 5% and 10% significance levels respectively. All figures are reported in basis points to two decimal places.

money markets and European Central Bank (ECB) monetary policy from 2007 onwards might bias the average *ex post* excess returns reported in table 2.8.

Eurozone money markets were immediately affected by the August 2007 money market turmoil.³⁶ Like the Federal Reserve, the ECB responded by injecting liquidity to money markets to preserve their proper functioning, which served to reduce the level of EONIA without adjusting the ECB's main refinancing (refi) rate.³⁷ On October 15, 2008, the ECB began to loosen monetary policy, reducing the refi rate from 4.25% to 3.75%. By May 13, 2009, the ECB refi rate had reached 1%.

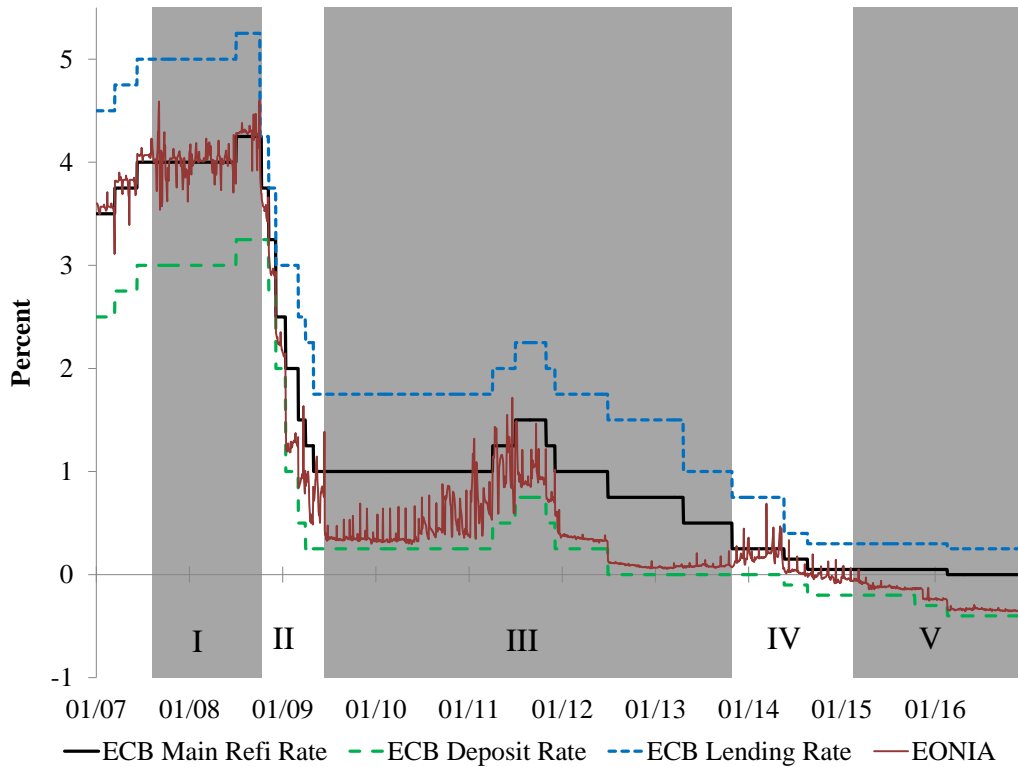
However, as a result of ECB liquidity operations, EONIA fell persistently below the refi rate from October 2008 onwards, as shown in figure 2.8. Between June 2009 and April 2011, the ECB's key interest rates (the refi rate, and the interest rates on the deposit and marginal lending facilities) were unchanged. EONIA was well below the refi rate during this period, closer to the ECB's standing deposit facility rate, indicating that the effective monetary policy stance was much looser than suggested by the main policy rate. Geraats (2011) labels this a period of 'monetary policy by stealth'. Although the ECB increased its refi rate twice in 2011, before reducing it five times between November 2011 and November 2013, EONIA continued to remain significantly below the headline policy rate. During this period the difference between the refi and deposit rates reduced from 75 basis points to 25 basis points, and the gap between EONIA and the refi rate narrowed, but widened again following cuts in the ECB deposit rate and an expanded asset purchase programme which was announced on January 22, 2015 and began on March 9, 2015. Because these developments significantly influenced the reference rate on Eurozone OIS contracts, I account for these periods in sensitivity analysis.

Table 2.9 provides summary statistics for the differences between EONIA and the refi rate

³⁶The difference between EONIA and the ECB's main refinancing (refi) rate increased from 9 basis points on August 8, 2007 to 22 basis points on August 9, 2007.

³⁷By August 28, 2007, EONIA was 28 basis points below the refi rate.

Figure 2.8: Eurozone Money Market Turmoil



Note: This figure depicts the ECB's refi, deposit and lending rates, and EONIA at a daily frequency from January 2, 2007 to December 31, 2016. The refi rate refers to the minimum bid or fixed rate for ECB main refinancing operations. The deposit rate refers to the rate on the ECB's deposit facility, which banks may use to make overnight deposits with the Eurosystem. The lending rate refers to the rate on the ECB's marginal lending facility, which offers overnight credit to banks from the Eurosystem. Area I denotes the initial period of money market turbulence, beginning on August 9, 2007 and ending on October 14, 2008. Area II represents the period of monetary policy loosening between October 15, 2008 and June 24, 2009. Area III denotes the period in which EONIA remained persistently below the refi rate, from June 25, 2009 to November 12, 2013. Area IV is an intermediate period between November 13, 2013 and March 8, 2015. Area V denotes the period in which EONIA remained persistently below the refi rate during the ECB's quantitative easing operations, from March 9, 2015 and December 31, 2016. *Data Source:* European Central Bank.

in five sub-samples from 2007 onwards. Prior to the money market turmoil — between January 2, 2007 and August 8, 2007 — EONIA was, on average, 5.8 basis points above the refi rate. During the period of money market turmoil — between August 9, 2007 and October 14, 2008 — the spread fell to 0.4 basis points on average, following ECB liquidity interventions that sought to stabilise EONIA around the refi rate. However, the spread was over twice as volatile.

From October 15, 2008, the ECB began to cut its refi rate, while also providing unlimited liquidity on demand through a fixed-tender procedure with full allotment at the refi rate. Between October 15, 2008 and June 24, 2009, EONIA was, on average, 40.8 basis points below the refi rate; the standard deviation of the difference was almost twice that in the period of initial money market turmoil. On June 24, 2009, the ECB initiated one-year longer-term refinancing operations. From June 25, 2009 to November 12, 2013, EONIA was, on average, 54.2 basis points below the refi rate, marking the first period of monetary policy by stealth. On November 13, 2013, the ECB's refi rate was cut by 25 basis points to 25 basis points. Between this date and March 8, 2015, the difference between EONIA and the ECB refi rate diminished — over the period, the average difference was just -7.6 basis points and the standard deviation of this

Table 2.9: Eurozone Money Market Turmoil: The Difference Between EONIA and the ECB Refi Rate in Percent Points

Dates	Mean	Standard Deviation	# Obs.
Pre-Turmoil (02/01/2007-08/08/2007)	0.058	0.063	154
I: Money Market Turmoil (09/08/2007-14/10/2008)	0.004	0.137	303
II: Monetary Policy Loosening (15/10/2008-24/06/2009)	-0.408	0.246	175
III: Monetary Policy By Stealth I (25/06/2009-12/11/2013)	-0.542	0.178	1128
IV: Intermediate Period (13/11/2013-08/03/2015)	-0.076	0.072	334
V: Monetary Policy By Stealth II (09/03/2015-31/12/2016)	-0.259	0.086	467

Note: Average difference between EONIA and the ECB refi rate in percent points using daily frequency data. The ECB refi rate is the minimum bid or fixed rate for main refinancing operations. The final column, “# Obs.” denotes the number of observations in each sub-sample. *Data Source:* European Central Bank.

difference was more comparable to that seen prior to the initial money market turmoil. However, when the ECB enacted its expanded asset purchase programme on March 9, 2015, EONIA fell further below the refi rate. Between March 9, 2015 and December 31, 2016, EONIA was, on average, 25.9 basis points below the refi rate.

I account for these five distinct periods in the sensitivity analysis indicated in figure 2.8 and table 2.9: (i) the initial money market turmoil beginning on August 9, 2007; (ii) the period of Eurozone monetary policy loosening, beginning on October 15, 2008; (iii) the initial period of monetary policy by stealth, beginning on June 25, 2009; (iv) the intermediate period, beginning on November 13, 2013; and (v) the second period of monetary policy by stealth, beginning on March 9, 2015 and running to the end of the sample. To do this, I augment the baseline regression with five dummy variables. As before, the dummy variables are set to unity on dates where the OIS contract horizon overlaps with the relevant period.

Table 2.10 presents the results of the augmented regression, where the estimated α_n^{OIS} coefficients can be interpreted as average *ex post* excess returns on Eurozone OIS contracts between January 3, 2000 and August 8, 2007.³⁸ The estimated dummy variable coefficients represent the increase in average *ex post* excess returns associated with the specific periods they pertain to. Importantly, the estimated α_n^{OIS} coefficients are insignificantly different from zero for the 1 to 3 and 7 to 24-month tenors, implying that these OIS contracts provide accurate information about investors’ expectations of future short-term interest rates. Although the 4 to 6-month coefficients are statistically significant, they are small (1.56 to 3.79 basis points).

The estimated dummy variable coefficients indicate that ECB monetary policy loosening between October 15, 2008 and June 6, 2009 had the largest positive impact on *ex post* excess returns. The estimated coefficients on this dummy variable are significantly positive at the 10% level, at least, for all tenors. This finding reflects the *ex ante* unexpected nature of the post-financial crisis monetary policy loosening, rather than risk premia within OIS contracts that undermine their use as measures of monetary policy expectations.

Interestingly, the coefficient on the first monetary policy by stealth dummy is significantly positive, at the 5% level at least, for the 1 to 3-month tenors. This indicates that the discrepancy between EONIA and the ECB’s refi rate that arose in these periods did have implications for the information contained in OIS rates, and their use as a measure of monetary policy expectations.

³⁸Because of the definition of the dummy variables, and the limited availability of 36-month Eurozone OIS rate data, I do not present estimates for this tenor in table 2.10.

Table 2.10: Average *Ex Post* Excess Returns on Eurozone OIS Contracts at Daily Frequency with Controls for Money Market Turmoil, Monetary Policy Loosening and Stealth

Maturity in Months	1	2	3	4	5	6
$\hat{\alpha}_n^{OIS}$	−0.14	0.12	0.64	1.56*	2.63**	3.79*
[<i>t</i> -statistic]	[−0.25]	[0.19]	[1.28]	[1.91]	[2.07]	[1.82]
Mon. Market Dummy	3.11*	4.69	5.63	6.20	6.51	6.82
[<i>t</i> -statistic]	[1.66]	[1.49]	[1.29]	[1.01]	[0.95]	[0.93]
Mon. Pol. Dummy	6.24*	17.19**	27.61***	37.64***	46.02***	52.79***
[<i>t</i> -statistic]	[1.76]	[2.25]	[2.66]	[3.95]	[8.52]	[7.99]
Policy by Stealth I Dummy	2.31**	2.62***	2.36***	1.51	0.70	0.43
[<i>t</i> -statistic]	[1.98]	[3.20]	[2.39]	[0.71]	[0.19]	[0.09]
Intermediate Dummy	−0.33	−0.66	−1.09	−1.47	−1.89	−2.58
[<i>t</i> -statistic]	[−0.40]	[−0.53]	[−0.66]	[−0.80]	[−0.96]	[−1.05]
Policy by Stealth II Dummy	0.05	−0.35	−0.66	−1.29	−1.93	−2.55
[<i>t</i> -statistic]	[0.10]	[−0.54]	[−0.96]	[−1.18]	[−1.35]	[−1.29]
Maturity in Months	7	8	9	10	11	12
$\hat{\alpha}_n^{OIS}$	4.83	5.52	5.51	4.64	4.24	3.89
[<i>t</i> -statistic]	[1.52]	[1.25]	[0.94]	[0.58]	[0.41]	[0.29]
Mon. Market Dummy	7.50	9.07	11.17	13.41	14.85	15.45
[<i>t</i> -statistic]	[0.90]	[0.91]	[0.87]	[0.79]	[0.72]	[0.66]
Mon. Pol. Dummy	58.09***	61.19***	61.86***	64.09***	65.41***	66.55***
[<i>t</i> -statistic]	[5.86]	[5.14]	[4.23]	[4.22]	[4.04]	[4.04]
Policy by Stealth I Dummy	0.70	2.22	6.01	11.41	16.34	21.62
[<i>t</i> -statistic]	[0.12]	[0.35]	[0.91]	[1.23]	[1.28]	[1.27]
Intermediate Dummy	−3.37	−4.39	−5.89*	−7.63***	−10.68***	−14.53***
[<i>t</i> -statistic]	[−1.15]	[−1.37]	[−1.83]	[−2.99]	[−4.18]	[−3.91]
Policy by Stealth II Dummy	−2.62	−1.90	−0.27	2.61	6.29	10.48
[<i>t</i> -statistic]	[−0.91]	[−0.47]	[−0.05]	[0.36]	[0.63]	[0.79]
Maturity in Months	15^a	18	21^a	24		
$\hat{\alpha}_n^{OIS}$	4.20	10.85	18.06	26.66		
[<i>t</i> -statistic]	[0.20]	[0.43]	[0.63]	[0.88]		
Mon. Market Dummy	9.47	−5.73	−20.14	−38.15		
[<i>t</i> -statistic]	[0.34]	[−0.20]	[−0.63]	[−1.01]		
Mon. Pol. Dummy	73.62***	87.67***	94.43***	98.84***		
[<i>t</i> -statistic]	[9.54]	[4.42]	[3.30]	[3.12]		
Policy by Stealth I Dummy	33.71	35.05	40.23	47.12		
[<i>t</i> -statistic]	[1.32]	[1.36]	[1.48]	[1.51]		
Intermediate Dummy	−27.19***	−36.06***	−48.18***	−58.59***		
[<i>t</i> -statistic]	[−2.63]	[−3.50]	[−12.75]	[−7.33]		
Policy by Stealth II Dummy	23.43	26.50	33.33	32.00*		
[<i>t</i> -statistic]	[1.00]	[1.07]	[1.58]	[1.68]		

Note: Results from regression (2.8) for Eurozone OIS contracts. Sample: Daily Frequency, Jan. 3, 2000 to Dec. 31, 2016, but for those indicated by ^a Aug. 22, 2001 to Dec. 31, 2016. Hodrick (1992) *t*-statistics in square brackets. 1%, 5% and 10% significance denoted by ***, * and *, respectively. All figures in basis points to two decimal places. ‘Mon. Market Dummy’: equal to 1 on dates where the OIS contract horizon overlaps with the Aug. 9, 2007 to Oct. 14, 2008 money market turmoil. ‘Mon. Pol. Dummy’: equal to 1 on dates where the OIS contract horizon overlaps with the Oct. 15, 2008 to Jun. 24, 2009 ECB monetary policy loosening. ‘Policy by Stealth I Dummy’ (‘Intermediate Dummy’): equal to 1 on dates where the OIS contract horizon overlaps with the Jun. 25, 2009 to Nov. 12, 2013 (Nov. 13, 2013 to Mar. 8, 2015) period. ‘Policy by Stealth II Dummy’: equal to 1 on dates where the OIS contract horizon overlaps with Mar. 9, 2015 to Dec. 31, 2016.

Table 2.11: Average *Ex Post* Excess Returns on Japanese OIS Contracts at Daily Frequency

	Panel A: Japan OIS Contracts				
Maturity in Months	1	2	3	4	5
$\hat{\alpha}_n^{OIS}$	-0.00	0.19	0.40	0.64	0.86
[<i>t</i> -statistic]	[-0.00]	[0.59]	[0.80]	[0.97]	[1.13]
Maturity in Months	6	9	12	18^a	24^a
$\hat{\alpha}_n^{OIS}$	1.28	2.61	4.12	8.04	12.94
[<i>t</i> -statistic]	[1.47]	[1.57]	[1.52]	[1.41]	[1.52]

Note: Results from regression (2.8) for Japanese OIS contracts. Sample: Daily Frequency, July 24, 2003 to December 31, 2016, but for those indicated by ^a December 7, 2005 to December 31, 2016 (due to OIS rate availability). Hodrick (1992) *t*-statistics are reported in square brackets. An excess return is significantly different from zero at the 1%, 5% and 10% significance level when the absolute value of the *t*-statistic exceeds 2.33, 1.96, 1.645 respectively. These are denoted with asterisks ***, ** and * for the 1%, 5% and 10% significance levels respectively. All figures are reported in basis points to two decimal places.

Overall, although events in the Eurozone call for a more nuanced study of *ex post* excess returns on OIS contracts, the above results indicate that, on average, 1 to 24-month tenors provide accurate measures of investors' interest rate expectations.

2.4.4 Japanese OIS Contracts

To calculate unconditional *ex post* excess returns on Japanese OIS contracts, I use TONAR as the overnight floating reference rate. The contracts have a two day spot lag and obey the *Actual 365* dating convention. I use Japanese OIS rates, between July 2003 and December 2016, of the following maturities: 1, 2, 3, 4, 5, 6, 9, 12, 18 and 24 months. ³⁹

Table 2.11 presents the estimated average *ex post* excess returns on Japanese OIS contracts. The most striking finding is that average *ex post* excess returns on Japanese contracts are much smaller quantitatively than those on US, UK and Eurozone contracts. This is, most likely, due to the smaller degree of variation in the Japanese policy rate during the 2003-2016 period, with Japanese policy bound by the ELB for most of this epoch. Average *ex post* excess returns are insignificantly different from zero for all maturities from 1 months to 2 years; these contracts accurately reflect investors' expectations of future short-term interest rates. The point estimate for the 1-month coefficient is zero, reflecting the relative constancy of Japanese policy rates over the sample period.

2.5 Conclusion

Three main results emerge from this chapter. First, and most importantly, 1 to 24-month US OIS rates, on average, provide accurate measures of investors' short-term interest rate expectations, an indicator of the *de facto* monetary policy stance. Average *ex post* excess returns on the majority of these contracts are insignificantly different from zero, and any significant results can be explained by unanticipated monetary policy loosening and money market turmoil in 2007-2008. These findings suggest that US OIS rates can be used as empirical measures of investors' expectations of future short-term interest rates out to 2 years in advance. Moreover, this result

³⁹The selection of maturities and horizon-specific sample periods are, again, determined by data availability. See appendix A.1.

supports the joint use of US OIS rates and the term structure of government bond yields to estimate longer-horizon monetary policy expectations in chapter 3.

Second, 1 to 11-month US OIS contracts provide measures of investors' interest rate expectations that are as good as those from comparable-horizon FFFs contracts. Excess returns on these OIS contracts are quantitatively similar to those on comparable-horizon portfolios of FFFs contracts.

Third, much of the accuracy of US OIS rates as financial market-based measures of future short-term interest rate expectations carries over to UK, Eurozone and Japanese OIS markets. OIS contracts with maturities of up to 2 years in the UK, Eurozone and Japanese OIS rates provide accurate measures of investors' interest rate expectations, with some exceptions for very short-maturity UK OIS contracts during the first half of the 2000s.

This has important implications for the understanding of monetary policy shocks on a global scale. To date, many methods used by monetary economists rely on FFF data to measure expectations of the future monetary policy stance (e.g. Gürkaynak et al., 2005a; Gertler and Karadi, 2015). This has limited the application of these methods to US data only. Motivated by the results in this chapter, researchers can look to OIS rates as a globally-comparable measure of monetary policy expectations that enables them to apply these methods to a wider set of countries. These results should serve as a useful reference for, *inter alia*, a developing literature on the global effects of domestic monetary policy shocks (e.g. Rey, 2016).

Chapter 3

Estimating Nominal Interest Rate Expectations: Overnight Indexed Swaps and the Term Structure

3.1 Introduction

Overnight indexed swap (OIS) rates tend to provide accurate information about the evolution of interest rate expectations out to a 2-year horizon. However, financial market participants and policymakers monitor a range of financial market instruments to attain real-time measures of longer-horizon expectations. Dynamic term structure models have increasingly been used to estimate and separately identify the evolution of expected future short-term interest rates and term premia (e.g. Gagnon, Raskin, Remache, and Sack, 2011; Christensen and Rudebusch, 2012, and chapter 4),¹ two components of nominal government bond yields. By imposing no-arbitrage, these models provide estimates of interest rate expectations that are consistent across the term structure, and extend to horizons in excess of what can be accurately imputed from financial market prices directly. However, a popular class of these models — Gaussian affine dynamic term structure models (GADTSMs) — suffers from an identification problem that results in estimates of interest rate expectations that are spuriously stable (e.g. Bauer, Rudebusch, and Wu, 2012; Kim and Orphanides, 2012; Guimarães, 2014).

Central to the identification problem is an informational insufficiency. Bond yield data is the sole input to an unaugmented GADTSM. These yields provide information of direct relevance to the estimation of the fitted bond yields. Absent additional information, estimates of interest rate expectations are poorly identified as they must also be derived from information contained within the actual bond yields. To do this, maximum likelihood or ordinary least squares estimates of, *inter alia*, the persistence of the (pricing factors derived from the) actual yields must be attained. However, as a symptom of the identification problem, a ‘finite sample’ bias will arise in these persistence parameters when there is insufficient information and a limited number of interest rate cycles in the observed yield data.² Finite sample bias will result in persistence parameters that are spuriously estimated to be less persistent than they really are and estimates

¹The term premium represents the compensation investors receive for, *inter alia*, default risk, interest rate risk and illiquidity.

²Kim and Orphanides (2012, p. 242) state that “in a term structure sample spanning 5 to 15 years, one may not observe a sufficient number of ‘mean reversion’.”

of future short-term interest rates that are spuriously stable.³ Because bond yields are highly persistent, the finite sample bias can be severe. Moreover, the severity of the bias is increasing in the persistence of the actual yield data. For daily frequency yields, which display greater persistence than lower-frequency data, the problem is particularly pertinent.

In this chapter, I propose the augmentation of GADTSMs with OIS rates as an additional estimation input to improve the identification of interest rate expectations and term premia from yields, building on the results in chapter 2. By using information to separately identify interest rate expectations, OIS-augmentation does tackle the informational insufficiency at the center of the GADTSM identification problem. To set up the OIS-augmented GADTSM, I derive expressions for the OIS pricing factor loadings that explicitly account for the payoff structure in OIS contracts. I estimate the OIS-augmented model using maximum likelihood via the Kalman filter with 3 to 24-month OIS rates, tenors at which OIS rates provide accurate information about interest rate expectations, and 3-month to 10-year US Treasury yields. The model provides estimates of interest rate expectations and term premia out to a 10-year horizon. To the extent that excess returns on OIS rates can vary on a day-to-day basis, I admit measurement error in the OIS excess returns over time in my OIS-augmented GADTSM. The Kalman filter maximum likelihood setup is well suited to account for this.

This is not the first proposed solution to the GADTSM identification problem. Kim and Orphanides (2012) suggest augmenting GADTSMs with survey expectations of future short-term interest rates for the same purpose. They document that, between 1990 and 2003, their survey-augmented model produces sensible estimates of interest rate expectations. Guimarães (2014) shows that, relative to an unaugmented GADTSM, the survey-augmented model provides interest rate expectation estimates that better correspond with survey expectations of future interest rates and delivers gains in the precision of interest rate expectation estimates. I show that estimated interest rate expectations from the OIS-augmented model are superior to the survey-augmented model for the 2002-2016 period.

Bauer et al. (2012) propose an alternative solution, focused on resolving the finite sample bias via bias-correction. They document that their bias-corrected estimates of interest rate expectations “are more plausible from a macro-finance perspective” (p. 454) than those from an unaugmented GADTSM. However, as Wright (2014) states, the fact that bias-correction has notable effects on GADTSM-estimated interest rate expectations is merely a *symptom* of the identification problem. Bias-correction does not directly address the identification problem at the heart of GADTSM estimation: the informational insufficiency. Moreover, Wright (2014) argues that the bias-corrected estimates of future interest rate expectations are “far too volatile” (p. 339). I find that estimated interest rate expectations from the OIS-augmented model are superior to bias-corrected estimates for the 2002-2016 period.

OIS-augmentation is closest in philosophy to survey-augmentation. The GADTSM is augmented with additional information to better identify the evolution of interest rate expectations. However, OIS-augmentation differs in a number of important respects, which help to explain its superior performance *vis-à-vis* survey-augmentation. Primarily, although survey forecasts do help to address the informational insufficiency problem, they are ill-equipped for the estimation

³This ‘finite sample’ bias is well documented for ordinary least squares estimation of a univariate autoregressive process, where estimates of the autoregressive parameter will be biased downwards, implying less persistence than the true process (Hurwicz, 1950). Within GADTSMs, the finite sample bias is a multivariate generalisation of this.

of daily frequency expectations. Survey forecasts of future interest rates are only available at a low frequency: quarterly or monthly, at best. Thus, survey forecasts are unlikely to provide sufficient information to accurately identify the daily frequency evolution of interest rate expectations. Moreover, the survey forecasts used by Kim and Orphanides (2012) and Guimarães (2014) correspond to the expectations of professional forecasters and not necessarily those of financial market participants.

OIS rates offer significant advantages over survey expectations for the daily frequency estimation of GADTSMs. Most importantly, OIS rates are available at a daily frequency, so provide information at the same frequency at which interest rate expectations are estimated. Secondly, OIS rates are formed as a result of actions by financial market participants, so can be expected to better reflect their expectations of future short-term interest rates. Third, the information in survey forecasts is limited in comparison to the expectational information contained in OIS rates. Survey forecasts typically provide information about expected future short-term interest rates for a short time period in the future.⁴ In contrast, there exists a term structure of OIS contracts that can be used to infer the evolution of investors' interest rate expectations from now until a specified future date. The horizon of these OIS contracts corresponds exactly to the horizon of nominal government bonds.

Away from the GADTSM-literature, OIS rates are increasingly being used to infer investors' expectations of future monetary policy, as chapter 2 discusses. Importantly, the OIS-augmented GADTSM is able to provide accurate measures of interest rate expectations, as well as term premia, at horizons in excess of 2 years. Moreover, the global comparability of OIS rates shown in chapter 2 implies that the OIS-augmented method is applicable in other advanced economies (see chapter 5).

OIS rates are also better suited to GADTSM-augmentation than federal funds futures rates, as the horizon of the former corresponds exactly to the horizon of the bond yield data used in GADTSMs. The horizon of a federal funds futures contract is a single month in the future, beginning on the first and ending on the last day of a specified month. Thus, OIS contracts provide a richer source of information with which to identify expected future short-term interest rates along the term structure.

I document that the OIS-augmented model accurately captures investors' expectations of future short-term interest rates out to the 10-year horizon. The in-sample model estimates of interest rate expectations co-move closely with federal funds futures rates and survey expectations of future short-term interest rates at horizons where comparison is possible. In these dimensions, the OIS-augmented model is superior to three other GADTSMs: (i) the unaugmented model, which only uses bond yield data to estimate both actual yields and interest rate expectations; (ii) the bias-corrected model of Bauer et al. (2012); and (iii) the survey-augmented model.⁵ The OIS-augmented model also best captures qualitative daily frequency movements in interest rate expectations implied by financial market instruments. Moreover, unlike other

⁴For example, the Federal Reserve Bank of Philadelphia's *Survey of Professional Forecasters* provides expectations of the average 3-month T-Bill rate during the current quarter, and the first, second, third and fourth quarters ahead.

⁵For the most direct comparison to the OIS-augmented model, I estimate the survey-augmented model using the algorithm of Guimarães (2014) which uses the same Joslin, Singleton, and Zhu (2011) identification restrictions as the OIS-augmented model, as opposed to the Kim and Wright (2005) survey-augmented model that applies the Kim and Orphanides (2012) identification algorithm, first proposed in Kim and Orphanides (2005). Chapter 4 shows that the Kim and Wright (2005) model performs worse than the OIS-augmented decomposition.

models, the interest rate expectations implied by the OIS-augmented model obey the zero lower bound for the US, despite the fact that additional restrictions are not imposed to achieve this. This represents an important contribution in the light of recent computationally burdensome proposals for term structure modelling at the zero lower bound (e.g. Christensen and Rudebusch, 2013a,b).

The remainder of this chapter is structured as follows. Section 3.2 describes OIS contracts and the accuracy of OIS-implied interest rate expectations. Section 3.3 lays out the unaugmented arbitrage-free GADTSM, before describing the identification problem and ‘finite sample’ bias with direct reference to the model parameters. Section 3.4 presents the OIS-augmented model. Section 3.5 documents the data and estimation methodology. Section 3.6 presents the results, documenting the superiority of the OIS-augmented model as a measure of interest rate expectations. Section 3.7 concludes.

3.2 Overnight Indexed Swaps

Chapter 2 details the features of overnight index swap (OIS) contracts, explains why changes in OIS rates can reasonably be associated with changes in investors’ expectations of future overnight interest rates over the horizon of the contract, and presents estimates of average *ex post* realised excess returns on OIS contracts. When accounting for 2007-2008 money market turmoil and the US monetary policy loosening of 2008 that was unexpected *ex ante*, the average *ex post* excess returns on 1 to 24-month US OIS contracts are statistically insignificant. 1 to 24-month OIS rates provide accurate measures of investors’ interest rate expectations, conforming to the expectations hypothesis (2.5), and verifying an important identifying assumption of the OIS-augmented GADTSM.

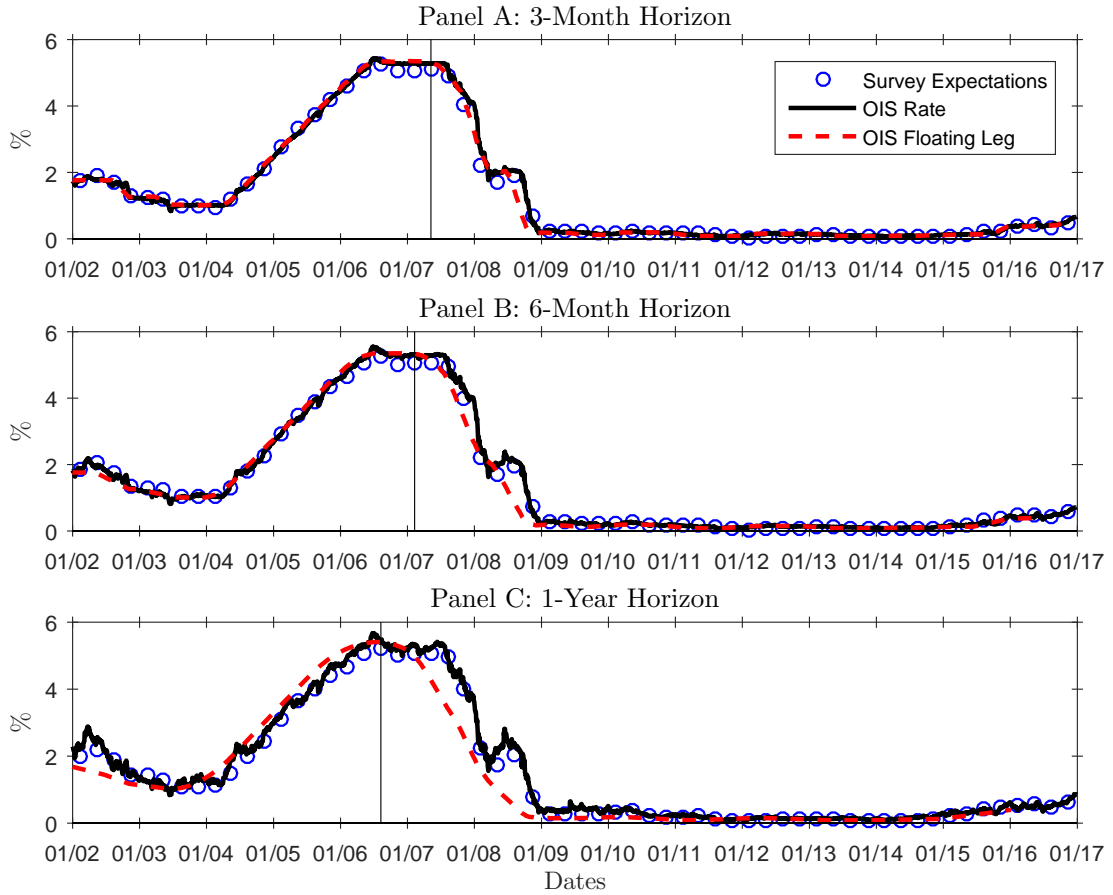
To further illustrate that short-horizon OIS rates provide accurate information about expectations of the future short-term interest rates, I compare OIS rates with survey expectations.

Figure 3.1 plots daily 3, 6 and 12-month OIS rates between January 2002 and December 2016 against both the daily frequency *ex post* realised floating leg of the swap and the quarterly frequency survey expectations of the future short-term nominal interest rate over the corresponding-horizon. I construct approximations of survey forecasts for the average 3-month US T-Bill rate using data from the *Survey of Professional Forecasters* (SPF) at the Federal Reserve Bank of Philadelphia.⁶ The survey is published every quarter and reports the median forecasters’ expectations of the average 3-month T-Bill rate over specified time periods: the current quarter $\bar{i}_{t|t}^{3m,sur}$; and the first $\bar{i}_{t+1|t}^{3m,sur}$, second $\bar{i}_{t+2|t}^{3m,sur}$, third $\bar{i}_{t+3|t}^{3m,sur}$ and fourth $\bar{i}_{t+4|t}^{3m,sur}$ quarters subsequent to the current one, where t denotes the current quarter. To construct the survey forecast approximations in figure 3.1, I first calculate the implied expectations of the average 3-month T-Bill rate over the remainder of the current quarter using the realised 3-month T-Bill rate over the current quarter up to the survey submission deadline date and the median survey expectation for the average 3-month T-Bill rate for the current quarter $\bar{i}_{t|t}^{3m,sur}$, exploiting the fact that the survey deadline dates lie approximately halfway through the ‘current’ quarter.⁷ Using this and the longer-horizon survey expectations, I then calculate geometric weighted averages of survey forecasts from the SPF (see appendix A.2). I use a geometric weighting scheme

⁶See Appendix A.1 for a detailed specification of data sources. Guimarães (2014) uses survey forecasts from the *Survey of Professional Forecasters* to decompose US interest rates.

⁷For example, the deadline date for the 2013 Q1 survey was February 11th 2013.

Figure 3.1: US OIS Rates and Corresponding *Ex Post* Realised Floating Rates, and Survey Expectations



Note: Daily OIS rates from *Bloomberg*. Daily *ex post* realised floating legs of the swaps calculated using equation (2.3). Survey expectations are from the *Survey of Professional Forecasters*. January 2002 to December 2016. The survey forecast, at each horizon, is attained by constructing the geometric weighted average of the median response of forecasters relating to their expectation of the average 3-month T-Bill rate over the relevant periods (see appendix A.2). Survey expectations are plotted on the forecast submission deadline date for each quarter. See appendix A.1 for detailed data source information. Vertical lines in each panel are plotted 3, 6 and 12 months prior to August 9, 2007 respectively, the date BNP Paribas froze funds citing US sub-prime mortgage sector problems.

to allow comparison with the geometric payoff structure of OIS contracts.⁸

Figure 3.1 plots survey expectations on submission dates, and demonstrates that survey and OIS-implied interest rate expectations co-moved closely between 2002 and 2016. The difference between the OIS rate i_t^{OIS} (solid black line) and the *ex post* realised floating leg i_t^{FLT} (dashed red line) graphically depicts the excess return defined in (2.4). Visual inspection of figure 3.1 confirms the formal results from chapter 2: OIS rates closely co-move with the *ex post* realised floating leg of the contracts. The most notable deviation of the two quantities occurs in 2007-2008, coinciding with the financial turmoil that erupted in this period. As the 2008 Federal Reserve policy easing was *ex ante* unanticipated, there is no reason to expect it to be reflected in *ex ante* expectations of future interest rates, explaining the difference in the quantities at this time. Similarly, there is a small difference between the 1-year OIS rate and the realised floating

⁸There are two caveats to this comparison which help to explain small differences between survey expectations and OIS rates. First, the expectational horizons of OIS rates and the T-Bill expectations do not exactly correspond, because the latter also reflect 3-month T-Bill rate expectations 1.5 months beyond the horizon, which reflect expected developments up to 4.5 months beyond the horizon. Second, 3-month T-Bill rates are on a discount basis, whereas OIS rates include expectations of interest rates on a yield basis.

rate during 2002, an artefact of unexpected US monetary policy loosening in response to the 2001 recession.

3.3 Term Structure Model

This section presents the discrete-time GADTSM that is commonplace in the literature (e.g. Ang and Piazzesi, 2003; Kim and Wright, 2005) and describes the identification problem in unaugmented GADTSMs with reference to model parameters. Since the focus of this chapter is on the identification of interest rate expectations and term premia at a daily frequency, hereafter t is a *daily* time index.⁹

3.3.1 Unaugmented Model Specification

The discrete-time GADTSM has three key foundations. First, there are K pricing factors \mathbf{x}_t (a $K \times 1$ vector), which follow a first-order vector autoregressive process under the actual probability measure \mathbb{P} :

$$\mathbf{x}_{t+1} = \boldsymbol{\mu} + \boldsymbol{\Phi}\mathbf{x}_t + \boldsymbol{\Sigma}\boldsymbol{\varepsilon}_{t+1} \quad (3.1)$$

where $\boldsymbol{\varepsilon}_{t+1}$ is a stochastic disturbance with the conditional distribution $\boldsymbol{\varepsilon}_{t+1}|\mathbf{x}_t \sim \mathcal{N}(\mathbf{0}_K, \mathbf{I}_K)$; $\mathbf{0}_K$ is a $K \times 1$ vector of zeros; and \mathbf{I}_K is a $K \times K$ identity matrix. $\boldsymbol{\mu}$ is a $K \times 1$ vector and $\boldsymbol{\Phi}$ is a $K \times K$ matrix of parameters. $\boldsymbol{\Sigma}$ is a $K \times K$ lower triangular matrix, which is invariant to the probability measure.

Second, the one-period short-term nominal interest rate i_t is assumed to be an affine function of the pricing factors:

$$i_t = \delta_0 + \boldsymbol{\delta}'_1 \mathbf{x}_t \quad (3.2)$$

where δ_0 is a scalar and $\boldsymbol{\delta}_1$ is a $K \times 1$ vector of parameters.

Third, no-arbitrage is imposed. The pricing kernel M_{t+1} that prices all assets when there is no-arbitrage is of the following form:

$$M_{t+1} = \exp\left(-i_t - \frac{1}{2}\boldsymbol{\lambda}'_t \boldsymbol{\lambda}_t - \boldsymbol{\lambda}'_t \boldsymbol{\varepsilon}_{t+1}\right) \quad (3.3)$$

where $\boldsymbol{\lambda}_t$ represents a $K \times 1$ vector of time-varying market prices of risk, which are affine in the pricing factors, following Duffee (2002):

$$\boldsymbol{\lambda}_t = \boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t \quad (3.4)$$

where $\boldsymbol{\lambda}_0$ is a $K \times 1$ vector and $\boldsymbol{\Lambda}_1$ is a $K \times K$ matrix of parameters.

The assumption of no-arbitrage guarantees the existence of a risk-adjusted probability measure \mathbb{Q} , under which the bonds are priced (Harrison and Kreps, 1979).¹⁰ Given the form of the market prices of risk in (3.4), the pricing factors \mathbf{x}_t also follow a first-order vector autoregressive

⁹The model can be estimated at lower frequencies, with the label for t changing correspondingly. Lloyd (2017a, Appendix F.3) presents a comparison of models estimated at a monthly frequency. The results from monthly frequency estimation are similar to those from daily frequency estimation.

¹⁰The risk-adjusted probability measure \mathbb{Q} is defined such that the price V_t of any asset that does not pay any dividends at time $t+1$ satisfies $V_t = \mathbb{E}_t^{\mathbb{Q}}[\exp(-i_t)V_{t+1}]$, where the expectation $\mathbb{E}_t^{\mathbb{Q}}$ is taken under the risk-adjusted probability measure \mathbb{Q} .

process under the risk-adjusted probability measure \mathbb{Q} :

$$\mathbf{x}_{t+1} = \boldsymbol{\mu}^{\mathbb{Q}} + \boldsymbol{\Phi}^{\mathbb{Q}} \mathbf{x}_t + \boldsymbol{\Sigma} \boldsymbol{\varepsilon}_{t+1}^{\mathbb{Q}} \quad (3.5)$$

where:¹¹

$$\boldsymbol{\mu}^{\mathbb{Q}} = \boldsymbol{\mu} - \boldsymbol{\Sigma} \boldsymbol{\lambda}_0, \quad \boldsymbol{\Phi}^{\mathbb{Q}} = \boldsymbol{\Phi} - \boldsymbol{\Sigma} \boldsymbol{\Lambda}_1.$$

and $\boldsymbol{\varepsilon}_{t+1}^{\mathbb{Q}}$ is a stochastic disturbance with the conditional distribution $\boldsymbol{\varepsilon}_{t+1}^{\mathbb{Q}} | \mathbf{x}_t \sim \mathcal{N}(\mathbf{0}_K, \mathbf{I}_K)$.

Bond Pricing Since M_{t+1} is the nominal pricing kernel that prices all nominal assets in the economy, the gross one-period return R_{t+1} on any nominal asset must satisfy:

$$\mathbb{E}_t [M_{t+1} R_{t+1}] = 1 \quad (3.6)$$

Let $P_{t,n}$ denote the price of an n -day zero-coupon bond at time t . Then, using $R_{t+1} = P_{t+1,n-1}/P_{t,n}$, (3.6) implies that the bond price is recursively defined:

$$P_{t,n} = \mathbb{E}_t [M_{t+1} P_{t+1,n-1}] \quad (3.7)$$

Alternatively, with no-arbitrage, the price of an n -period zero-coupon bond must also satisfy the following relation under the risk-adjusted probability measure \mathbb{Q} :

$$P_{t,n} = \mathbb{E}_t^{\mathbb{Q}} [\exp(-i_t) P_{t+1,n-1}] \quad (3.8)$$

By combining the dynamics of the pricing factors (3.5) and the short-term interest rate (3.2) with (12), the bond prices can be shown to be exponentially affine function in the pricing factors:

$$P_{t,n} = \exp(\mathcal{A}_n + \mathcal{B}_n \mathbf{x}_t) \quad (3.9)$$

where the scalar $\mathcal{A}_n \equiv \mathcal{A}_n(\delta_0, \boldsymbol{\delta}_1, \boldsymbol{\mu}^{\mathbb{Q}}, \boldsymbol{\Phi}^{\mathbb{Q}}, \boldsymbol{\Sigma}; \mathcal{A}_{n-1}, \mathcal{B}_{n-1})$ and $\mathcal{B}_n \equiv \mathcal{B}_n(\boldsymbol{\delta}_1, \boldsymbol{\Phi}^{\mathbb{Q}}; \mathcal{B}_{n-1})$, a $1 \times K$ vector, are recursively defined loadings:¹²

$$\begin{aligned} \mathcal{A}_n &= -\delta_0 + \mathcal{A}_{n-1} + \frac{1}{2} \mathcal{B}_{n-1} \boldsymbol{\Sigma} \boldsymbol{\Sigma}' \mathcal{B}_{n-1}' + \mathcal{B}_{n-1} \boldsymbol{\mu}^{\mathbb{Q}} \\ \mathcal{B}_n &= -\boldsymbol{\delta}_1' + \mathcal{B}_{n-1} \boldsymbol{\Phi}^{\mathbb{Q}} \end{aligned}$$

with initial values $\mathcal{A}_0 = 0$ and $\mathcal{B}_0 = \mathbf{0}'_K$ ensuring that the price of a ‘zero-period’ bond is one.

The continuously compounded yield on an n -day zero-coupon bond at time t , $y_{t,n} = -\frac{1}{n} \ln(P_{t,n})$, is given by:

$$y_{t,n} = A_n + B_n \mathbf{x}_t \quad (3.10)$$

where $A_n \equiv -\frac{1}{n} \mathcal{A}_n(\delta_0, \boldsymbol{\delta}_1, \boldsymbol{\mu}^{\mathbb{Q}}, \boldsymbol{\Phi}^{\mathbb{Q}}, \boldsymbol{\Sigma}; \mathcal{A}_{n-1}, \mathcal{B}_{n-1})$ and $B_n \equiv -\frac{1}{n} \mathcal{B}_n(\boldsymbol{\delta}_1, \boldsymbol{\Phi}^{\mathbb{Q}}; \mathcal{B}_{n-1})$.

The risk-neutral yield on an n -day bond reflects the expectation of the average short-term interest rate over the n -day life of the bond, corresponding to the yields that would prevail

¹¹See appendix B.1.2 for a formal derivation of these expressions.

¹²See appendix B.1.1 for a formal derivation of these expressions.

if investors were risk-neutral.¹³ That is, the yields that would arise under the expectations hypothesis of the yield curve. The risk-neutral yields can be calculated using:

$$\tilde{y}_{t,n} = \tilde{A}_n + \tilde{B}_n \mathbf{x}_t \quad (3.11)$$

where $\tilde{A}_n \equiv -\frac{1}{n} \mathcal{A}_n(\delta_0, \delta_1, \boldsymbol{\mu}, \boldsymbol{\Phi}, \boldsymbol{\Sigma}; \mathcal{A}_{n-1}, \mathcal{B}_{n-1})$ and $\tilde{B}_n \equiv -\frac{1}{n} \mathcal{B}_n(\delta_1, \boldsymbol{\Phi}; \mathcal{B}_{n-1})$.¹⁴ Note that, the risk-neutral yields are attained, *inter alia*, using parameters specific to the actual probability measure \mathbb{P} , $\{\boldsymbol{\mu}, \boldsymbol{\Phi}\}$. But, because no-arbitrage is assumed, the bonds are priced under the risk-adjusted measure \mathbb{Q} , so the fitted yields are attained, *inter alia*, by using parameters specific to the risk-adjusted probability measure \mathbb{Q} , $\{\boldsymbol{\mu}^{\mathbb{Q}}, \boldsymbol{\Phi}^{\mathbb{Q}}\}$.

The spot term premium on an n -day bond is defined as the difference between the fitted yield (3.10) and the risk-neutral yield (3.11):

$$tp_{t,n} = y_{t,n} - \tilde{y}_{t,n} \quad (3.12)$$

3.3.2 Unaugmented GADTSMs and the Identification Problem

Numerous studies have documented problems with separately identifying expectations of future short-term interest rates (risk-neutral yields) from term premia (e.g. Bauer et al., 2012; Kim and Orphanides, 2012; Guimarães, 2014). The underlying source of difficulty is an informational insufficiency, which gives rise to finite sample bias.

The unaugmented model uses zero-coupon bond yield data as its sole input. This data provides a complete set of information about the dynamic evolution of the cross-section of yields — the yield curve. This provides sufficient information to accurately identify the risk-adjusted \mathbb{Q} dynamics — specifically, the parameters $\{\boldsymbol{\mu}^{\mathbb{Q}}, \boldsymbol{\Phi}^{\mathbb{Q}}\}$ in (3.5) — which (3.10) shows are of direct relevance to estimating *actual* yields. However, if there is no additional information and the sample of yields contains too few interest rate cycles,¹⁵ this data is not sufficient for the identification of the actual \mathbb{P} dynamics — specifically, the parameters $\{\boldsymbol{\mu}, \boldsymbol{\Phi}\}$ in (3.1) — which (3.11) illustrates are of relevance to the estimation of *risk-neutral* yields.¹⁶ Estimates of $\boldsymbol{\Phi}$ for the autoregressive process in (3.1) will suffer from finite sample bias. In particular, the persistent yields will have persistent pricing factors, so maximum likelihood or ordinary least squares estimates of the persistence parameters of the vector autoregressive process in (3.1) $\boldsymbol{\Phi}$ will be biased downwards.¹⁷ That is, the estimated $\hat{\boldsymbol{\Phi}}$ will understate the true persistence of the pricing factors, implying a spuriously fast mean reversion of future short-term interest rates. Because, in the model, agents form expectations of future short-term interest rates based on estimates of pricing factor mean reversion in $\hat{\boldsymbol{\Phi}}$, their estimates of the future short-term interest rate path will mean revert spuriously quickly too. Consequently, the estimated risk-neutral yields, which summarise the average of the expected path of future short-term interest rates, will vary little and will not accurately reflect the evolution of interest rate expectations.

¹³There is a small difference between risk-neutral yields and expected yields due to a convexity effect. In the homoskedastic model considered here, these effects are constant for each maturity and, in practice, small, corresponding to the $\frac{1}{2} \mathcal{B}_{n-1} \boldsymbol{\Sigma} \boldsymbol{\Sigma}' \mathcal{B}_{n-1}'$ term in the recursive expression for \mathcal{B}_n above.

¹⁴See appendix B.1.3 for a formal derivation of these expressions.

¹⁵Kim and Orphanides (2012, p. 242) state that 5 to 15-year samples may contain too few interest rate cycles.

¹⁶Note that because $\boldsymbol{\mu} = \boldsymbol{\mu}^{\mathbb{Q}} + \boldsymbol{\Sigma} \boldsymbol{\lambda}_0$ and $\boldsymbol{\Phi} = \boldsymbol{\Phi}^{\mathbb{Q}} + \boldsymbol{\Sigma} \boldsymbol{\Lambda}_1$, estimates of the time-varying market prices of risk, $\boldsymbol{\lambda}_0$ and $\boldsymbol{\Lambda}_1$, are required to estimate $\{\boldsymbol{\mu}, \boldsymbol{\Phi}\}$ and the risk-neutral yields.

¹⁷This is a multivariate generalisation of the downward bias in the estimation of autoregressive parameters by ordinary least squares in the univariate case.

The magnitude of the finite sample bias is increasing in the persistence of the data. For daily frequency yield data, which is highly persistent, the bias will be more severe. This not only motivates the augmentation of the GADTSM with additional data, but motivates the use of additional *daily frequency* data, namely OIS rates.

3.4 The OIS-Augmented Model

I estimate the OIS-augmented model using Kalman filter-based maximum likelihood. The Kalman filtering approach is particularly convenient for the augmentation of GADTSMs, as it can handle mixed-frequency data. Specifically, for OIS-augmentation, this allows estimation of the GADTSM for periods extending beyond that for which OIS rates are available.¹⁸

To implement the Kalman filter-based estimation, I use (3.1), the vector autoregression for the latent pricing factors under the actual probability measure \mathbb{P} , as the transition equation.

The observation equation depends on whether OIS rates are observed on day t . On days when the OIS rates are *not* observed (i.e. days prior to January 2002), the observation equation is formed by stacking the N yield maturities in (3.10) to form:

$$\mathbf{y}_t = \mathbf{A} + \mathbf{B}\mathbf{x}_t + \Sigma_Y \mathbf{u}_t \quad (3.13)$$

where: $\mathbf{y}_t = [y_{t,n_1}, \dots, y_{t,n_N}]'$ is the $N \times 1$ vector of bond yields; $\mathbf{A} = [A_{n_1}, \dots, A_{n_N}]'$ is an $N \times 1$ vector and $\mathbf{B} = [B'_{n_1}, \dots, B'_{n_N}]'$ is an $N \times K$ matrix of bond-specific loadings; $A_{n_\iota} = -\frac{1}{n_\iota} \mathcal{A}_{n_\iota}(\delta_0, \delta_1, \mu^Q, \Phi^Q, \Sigma; \mathcal{A}_{n_\iota-1}, \mathcal{B}_{n_\iota-1})$ and $B_{n_\iota} = -\frac{1}{n_\iota} \mathcal{B}_{n_\iota}(\delta_1, \Phi^Q; \mathcal{B}_{n_\iota-1})$ are the bond-specific loadings; and $\iota = 1, 2, \dots, N$ such that n_ι denotes the maturity of bond ι in days. The $N \times 1$ vector $\mathbf{u}_t \sim \mathcal{N}(\mathbf{0}_N, \mathbf{I}_N)$ denotes the yield measurement error, where $\mathbf{0}_N$ is an N -vector of zeros and \mathbf{I}_N is an $N \times N$ identity matrix. Here, like much of the existing literature,¹⁹ I impose a homoskedastic form for the yield measurement error, such that Σ_Y is an $N \times N$ diagonal matrix with common diagonal element σ_e , the standard deviation of the yield measurement error. The homoskedastic error is characterised by a single parameter σ_e , maintaining computational feasibility for an already high-dimensional maximum likelihood routine.

On days when OIS rates are observed, the Kalman filter observation equation is augmented with OIS rates. The following proposition illustrates that OIS rates can (approximately) be written as an affine function of the pricing factors with loadings A_j^{ois} and B_j^{ois} for J different OIS maturities, where $j = j_1, j_2, \dots, j_J$ denote the J OIS horizons in days. The loadings presented in this proposition are calculated by assuming that the expectations hypothesis (2.5) holds for the OIS tenors included in the model, an assumption that was verified in section 3.2 for the maturities used here. Moreover, the loadings explicitly account for the payoff structure of an OIS contract. It is in this respect that the technical setup of the OIS-augmented GADTSM most clearly differs from the survey-augmented model.

¹⁸This thesis uses daily US OIS rates from 2002, the first date for which these rates are consistently available at all the relevant tenors on Bloomberg. Models are estimated from this date to directly isolate the effect of OIS rates on GADTSMs. However, given the Kalman filter method, the model can be estimated over longer periods.

¹⁹See, for example, Guimarães (2014).

Proposition The j -day OIS rate on date t $i_{t,t+j}^{ois}$, where $j = j_1, j_2, \dots, j_J$, can be (approximately) written as an affine function of the pricing factors \mathbf{x}_t :

$$i_{t,t+j}^{ois} = A_j^{ois} + B_j^{ois} \mathbf{x}_t \quad (3.14)$$

where $A_j^{ois} \equiv \frac{1}{j} \mathcal{A}_j^{ois} \left(\delta_0, \delta_1, \boldsymbol{\mu}, \boldsymbol{\Phi}, \boldsymbol{\Sigma}; \mathcal{A}_{j-1}^{ois}, \mathcal{B}_{j-1}^{ois} \right)$ and $B_j^{ois} \equiv \frac{1}{j} \mathcal{B}_j^{ois} \left(\delta_1, \boldsymbol{\Phi}; \mathcal{B}_{j-1}^{ois} \right)$ are recursively defined as:

$$\begin{aligned} \mathcal{A}_j^{ois} &= \delta_0 + \delta_1' \boldsymbol{\mu} + \mathcal{A}_{j-1}^{ois} + \mathcal{B}_{j-1}^{ois} \boldsymbol{\mu} \\ \mathcal{B}_j^{ois} &= \delta_1' \boldsymbol{\Phi} + \mathcal{B}_{j-1}^{ois} \boldsymbol{\Phi} \end{aligned}$$

where $\mathcal{A}_0^{ois} = 0$ and $\mathcal{B}_0^{ois} = \mathbf{0}_K'$, where $\mathbf{0}_K$ is a $K \times 1$ vector of zeros.

Proof: See appendix B.2.

Given this, the Kalman filter observation equation on the days OIS rates are observed is:

$$\begin{bmatrix} \mathbf{y}_t \\ \mathbf{i}_t^{ois} \end{bmatrix} = \begin{bmatrix} \mathbf{A} \\ \mathbf{A}^{ois} \end{bmatrix} + \begin{bmatrix} \mathbf{B} \\ \mathbf{B}^{ois} \end{bmatrix} \mathbf{x}_t + \begin{bmatrix} \boldsymbol{\Sigma}_Y & \mathbf{0}_{N \times J} \\ \mathbf{0}_{J \times N} & \boldsymbol{\Sigma}_O \end{bmatrix} \begin{bmatrix} \mathbf{u}_t \\ \mathbf{u}_t^{ois} \end{bmatrix} \quad (3.15)$$

where, in addition to the definitions of \mathbf{y}_t , \mathbf{A} , \mathbf{B} , $\boldsymbol{\Sigma}_Y$ and \mathbf{u}_t above, $\mathbf{i}_t^{ois} = [i_{t,j_1}^{ois}, \dots, i_{t,j_J}^{ois}]'$ is the $J \times 1$ vector of OIS rates; $\mathbf{A}^{ois} = [A_{j_1}^{ois}, \dots, A_{j_J}^{ois}]'$ is a $J \times 1$ vector and $\mathbf{B}^{ois} = [B_{j_1}^{ois}, \dots, B_{j_J}^{ois}]'$ is a $J \times K$ matrix of OIS-specific loadings; $\mathbf{0}_{N \times J}$ and $\mathbf{0}_{J \times N}$ denote $N \times J$ and $J \times N$ matrices of zeros respectively; and $\mathbf{u}_t^{ois} \sim \mathcal{N}(\mathbf{0}_J, \mathbf{I}_J)$ denotes the OIS measurement error, where $\mathbf{0}_J$ is an J -vector of zeros and \mathbf{I}_J is an $J \times J$ identity matrix. The inclusion of the measurement error permits non-zero OIS forecast errors, imposing that the forecast error is zero *on average*. I compared two parameterisations of $\boldsymbol{\Sigma}_O$; a homoskedastic model, with common diagonal elements in $\boldsymbol{\Sigma}_O$, and a heteroskedastic model, with distinct diagonal elements. A likelihood ratio test of the two did not reject the null hypothesis that all diagonal elements are equal, so I impose a homoskedastic form for the OIS measurement error such that $\boldsymbol{\Sigma}_O$ has common diagonal element σ_o , the standard deviation of the OIS measurement error, and zero elsewhere. The homoskedastic OIS measurement errors also provide computational benefits, as there are fewer parameters to estimate than if a more general covariance structure was permitted.²⁰

3.5 Methodology

To compare the OIS-augmented model with the existing literature, I estimate the following GADTSM-variants: (i) an unaugmented OLS/ML model, estimated using the Joslin et al. (2011) identification scheme, where K portfolios of yields are observed without error and are measured with the first K estimated principal components of the bond yields; (ii) the Bauer et al. (2012) bias-corrected model; (iii) a survey-augmented model, using expectations of future short-term interest rates for the subsequent four quarters as an additional input, estimated with

²⁰Kim and Orphanides (2012) and Guimarães (2014) impose homoskedasticity on the survey measurement errors in their Kalman filter setup for this reason.

the Kalman filter using the algorithm of Guimarães (2014) (see appendix B.3 for details);²¹ and (iv) the OIS-augmented model.

3.5.1 Data

In all models, Use the following bond yields \mathbf{y}_t : 3 and 6 months, 1 year, 18 months, 2 years, 30 months, 3 years, 42 months, 4 years, 54 months, 5, 7 and 10 years.²² For the 3 and 6-month yields, I use US T-Bill rates in accordance with much of the existing dynamic term structure literature.²³ The remaining rates are from the continuously compounded zero-coupon yields of Gürkaynak, Sack, and Wright (2007a). This data is constructed from daily-frequency fitted Nelson-Siegel-Svensson yield curves. Using the parameters of these curves, which are published along with the estimated zero-coupon yield curve, I back out the cross-section of yields for the 11 maturities from 1 to 10-years.

OIS rates are from *Bloomberg*. I use combinations of 3, 6, 12 and 24-month OIS rates in the OIS-augmented models. The choice of these maturities is motivated by evidence in section 3.2 and chapter 2. I estimate three variants of the OIS-augmented model. The first, baseline setup, includes the 3, 6, 12 and 24-month OIS rates (4-OIS-Augmented model). The second and third models include the 3, 6 and 12-month (3-OIS-Augmented model) *and* 3 and 6-month (2-OIS-Augmented model) tenors respectively.²⁴ Of the three OIS-augmented models, I find that the 4-OIS-Augmented model provides risk-neutral yields that best fit the evolution of interest rate expectations, in and out-of-sample.

Since US OIS rates are consistently available from January 2002, the baseline sample period runs from January 2002 to December 2016 to isolate the effect of OIS augmentation.²⁵

In accordance with the evidence of Litterman and Scheinkman (1991), that the first three principal components of bond yields explain well over 95% of their variation, I estimate the models with three pricing factors ($K = 3$).²⁶ By using the three-factor specification, for which the pricing factors have a well-understood economic meaning (the level, slope and curvature of the yield curve respectively), I am able to isolate and explain the economic mechanisms through which the OIS-augmented model provides superior estimates of expectations of future short-term interest rates *vis-à-vis* the unaugmented, bias-corrected and survey-augmented models.

²¹For direct comparison to the OIS-augmented model, I estimate the survey-augmented model with the Guimarães (2014) algorithm, which uses the same Joslin et al. (2011) identification scheme. Kim and Orphanides (2012) implement a different identification scheme in the estimation of their survey-augmented model. Like Guimarães (2014), I use survey expectations from the *Survey of Professional Forecasters* at the Federal Reserve Bank of Philadelphia, including forecasts of the 3-month T-Bill rate for the remainder of the current quarter and the first, second, third and fourth quarters ahead.

²²These yield maturities correspond to those used by Adrian, Crump, and Moench (2013).

²³The T-Bill rates are converted from their discount basis to the yield basis.

²⁴Because the results from the 2-OIS-augmented model are inferior to those from the 4 and 3-OIS-augmented models, I present results for the 2-OIS-augmented model in appendix B.4. The 2-OIS model may perform less well because the 3 and 6-month OIS rates that augment the model add little information on interest rate expectations over-and-above the 3 and 6-month T-Bill rates. The 3 and 4-OIS models benefit from longer-maturity OIS rates.

²⁵I estimate this linear model over this sample to make full use of OIS rate data. There may be benefits from combining OIS-augmentation with a non-linear model to fully account for the ELB.

²⁶I also estimate a four-factor specification in the light of evidence by Cochrane and Piazzesi (2005, 2008) and Duffee (2011) who argue that more than three factors are necessary to explain the evolution of nominal Treasury yields. These results are discussed in appendix B.4.2.

3.5.2 Estimation

The OIS-augmented model relies on Kalman filter-based maximum likelihood estimation, for which the pricing factors \mathbf{x}_t are latent. Normalisation restrictions must be imposed on the parameters to achieve identification. For this, I appeal to the Joslin et al. (2011) normalisation scheme, which allows “computationally efficient estimation of G[A]DTSMs” (Joslin et al., 2011, p. 928) and fosters faster convergence to the global optimum of the model’s likelihood function than other normalisation schemes (e.g. Dai and Singleton, 2000).²⁷ This permits a two-stage approach to estimating the OIS-augmented model.

To benefit fully from the computational efficiency of the Joslin et al. (2011) normalisation scheme, I first estimate the unaugmented GADTSM (hereafter, the OLS/ML model), presented in section 3.3.1, assuming that K portfolios of yields are priced without error, to attain initial values for the Kalman filter used in the second estimation stage. In particular, these K yield ‘portfolios’, \mathbf{x}_t , correspond to the first K estimated principal components of the bond yields. Under the Joslin et al. (2011) normalisation, this itself enables a two sub-stage estimation: first the \mathbb{P} parameters are estimated by OLS on equation (3.1) using the K estimated principal components in the vector \mathbf{x}_t ; second the \mathbb{Q} parameters are estimated by maximum likelihood (see appendix B.3).

Having attained these OLS/ML parameter estimates, I estimate the OIS-augmented model — which assumes all yields are observed with error — using the OLS/ML parameter estimates as initial values for the Kalman filter-based maximum likelihood routine.

3.6 Term Structure Results

3.6.1 Model Fit

This sub-section discusses four aspects of model fit: estimated bond yields, OIS rates, pricing factors and parameters.

Fitted Bond Yields

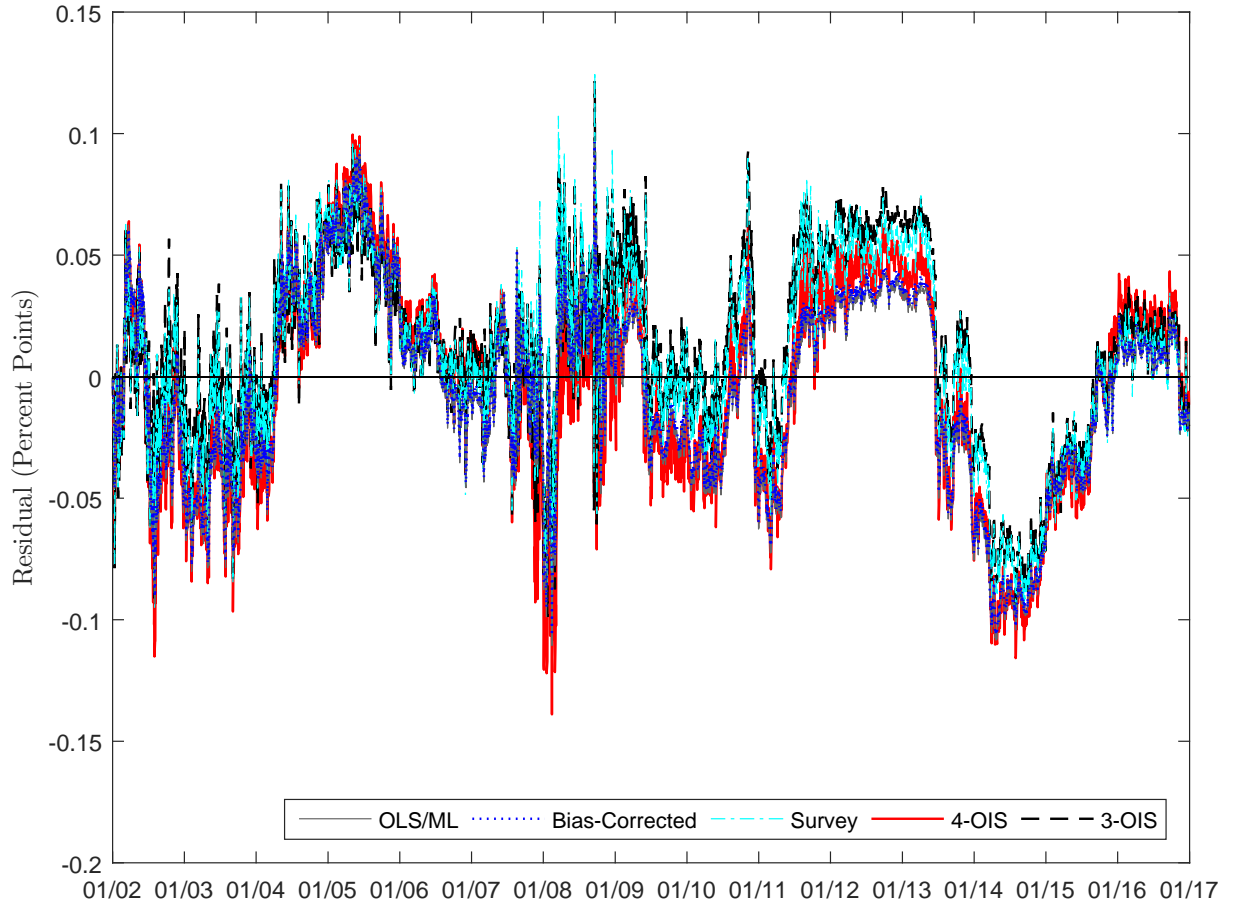
Importantly, OIS-augmentation of GADTSMs does not compromise the overall model fit with respect to actual bond yields. The fit of actual yields is strikingly similar across all the models. Figure 3.2 illustrates that the residuals of the 2-year fitted yield from the OLS/ML, bias-corrected, survey-augmented, 4 and 3-OIS-augmented models follow similar qualitative and quantitative paths.²⁸

The similar fit of actual yields is intuitive. I augment the GADTSM with OIS rates to provide additional information with which to better estimate parameters under the actual probability measure $\mathbb{P} \{\boldsymbol{\mu}, \boldsymbol{\Phi}\}$, which directly influence estimates of the risk-neutral yields. Estimates of

²⁷The computational benefits of the Joslin et al. (2011) normalisation scheme arise because it only imposes restrictions on the short-term interest rate i_t and the factors \mathbf{x}_t under the \mathbb{Q} probability measure. Consequently, the \mathbb{P} and \mathbb{Q} dynamics of the model do not exhibit strong dependence. Under the Dai and Singleton (2000) scheme, restrictions on the volatility matrix $\boldsymbol{\Sigma}$, which influences both the \mathbb{P} and \mathbb{Q} evolution of the factors (see equations (3.1) and (3.5)), create a strong dependence between the parameters under the two probability measures, engendering greater computational complexity in the estimation.

²⁸Table B.1, in appendix B.4.1, provides more detailed evidence of the similar actual yield fit of the models, documenting the root mean square error (RMSE) for each model at each yield maturity. Over the 13 maturities, the average RMSE of each model is around 5 basis points.

Figure 3.2: Residual of the 2-Year Fitted Yield from GADTSMs



Note: Residuals of the 2-year fitted yield from five GADTSMs: (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML); (ii) the bias-corrected model (Bias-Corrected); (iii) the survey-augmented model (Survey); (iv) the 4-OIS-augmented model (4-OIS); and (v) the 3-OIS-augmented model (3-OIS). The models are estimated with three pricing factors, using daily data from January 2002 to December 2016. The residual is defined as the actual yield subtracted by the model-implied fitted yield. The residuals are presented in annualised percentage points.

the fitted yield depend upon the risk-adjusted measure \mathbb{Q} parameters $\{\mu^{\mathbb{Q}}, \Phi^{\mathbb{Q}}\}$, which are not directly influenced by the OIS rates in the model, and are well-identified with bond yield data that provide information on the dynamic evolution of the cross-section of yields.

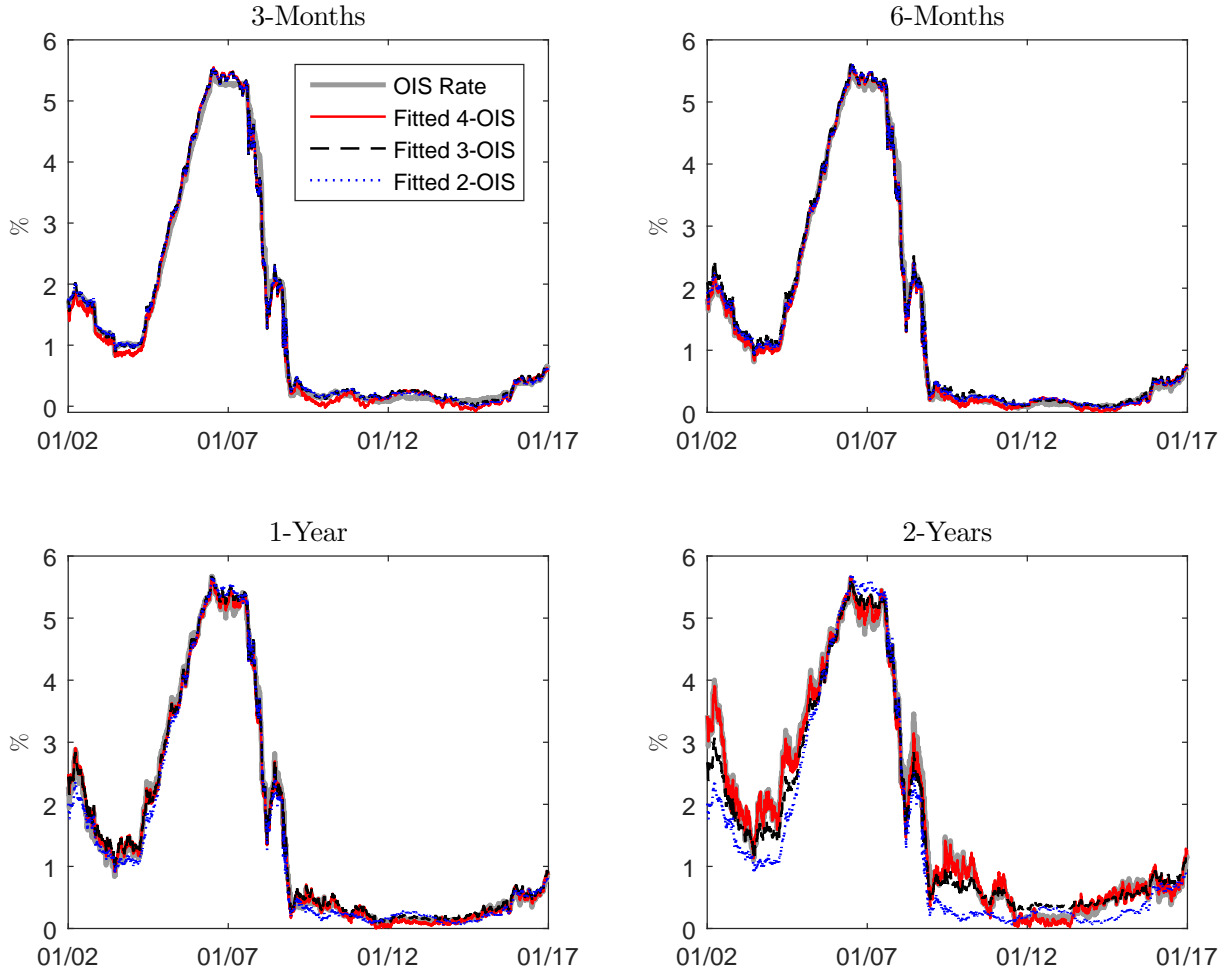
Fitted OIS Rates

The OIS-augmented models also provide fitted values for OIS rates. Figure 3.3 plots the 3, 6, 12 and 24-month OIS rates alongside the corresponding-maturity fitted-OIS rates from the 4, 3 and 2-OIS-augmented models.

The plots illustrate that the OIS-augmented models provide accurate estimates of actual OIS rates.²⁹ The 4-OIS-augmented model provides the best fit for the 6, 12 and 24-month OIS rates, while the 2-OIS-augmented model best fits the 3-month OIS rate. Although the differences between the OIS-augmented models at the 3-month horizon are marginal, the 4-OIS-augmented model fits the 24-month OIS rate substantially better than the 3 and 2-OIS-augmented models. This is unsurprising, as this OIS tenor is observed in the 4-OIS-augmented model. The 2-OIS-augmented model fits the 1 and 2-year OIS rates least well. This is unsurprising, as it uses the

²⁹Table B.2, in appendix B.4.1, provides detailed numerical evidence.

Figure 3.3: Fitted OIS Rates from the OIS-Augmented Models



Note: Fitted and actual 3, 6, 12 and 24-month OIS rates. Fitted OIS rates are from the 4, 3 and 2-OIS-augmented GADTSMs. The models are estimated with three pricing factors using daily data from January 2002 to December 2016. All figures are in annualised percentage points.

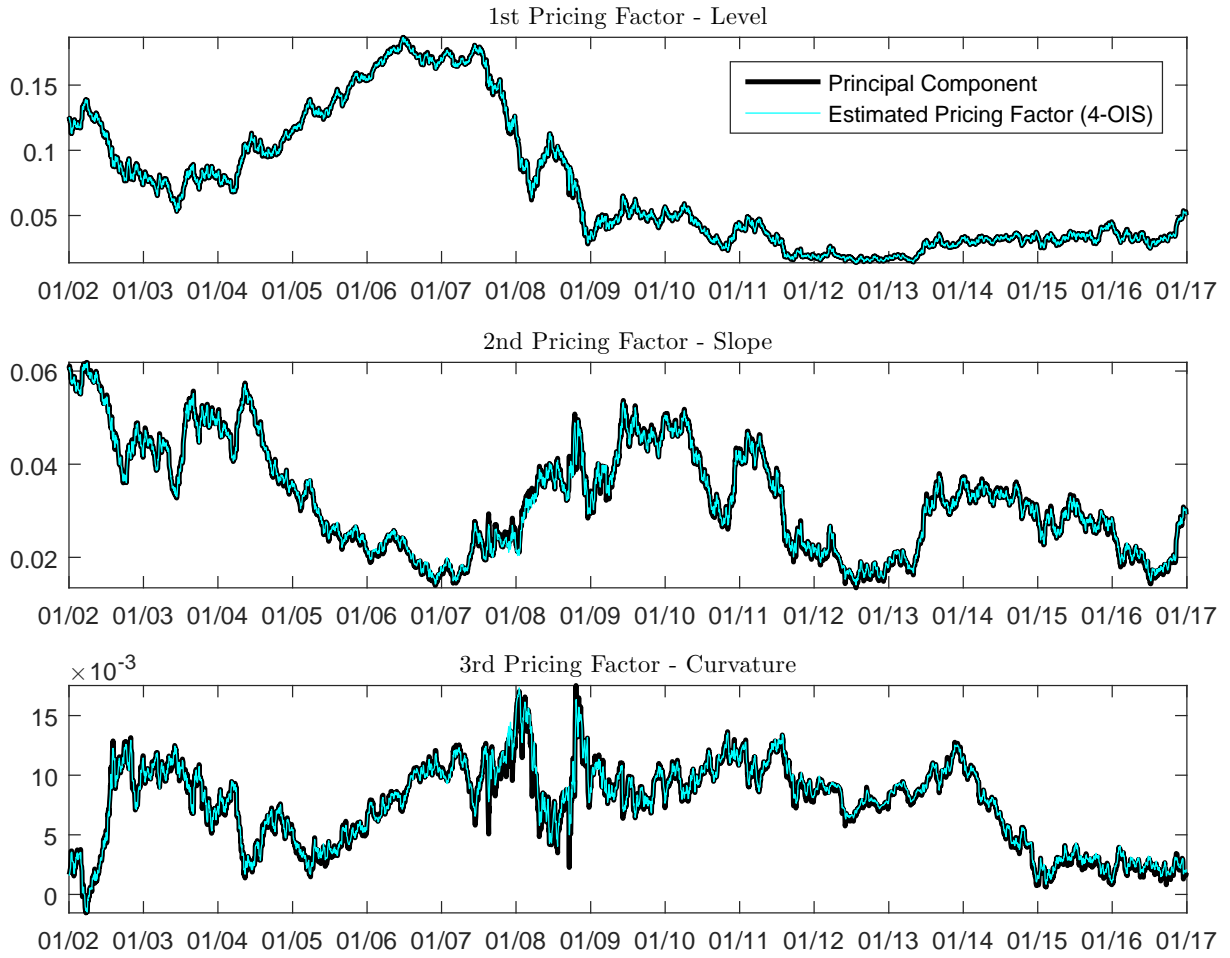
fewest OIS rates as observable inputs.

The fact the OIS-augmented models do not fit OIS rates as well as they fit bond yields — the quantitative value of OIS-RMSE (approximately 10 basis points) is almost double that of the bond yield-RMSE (approximately 5 basis points) — is neither worrying nor surprising. The GADTSM uses thirteen bond yields as inputs to estimate the cross-section of fitted yields in every time period, whereas only four OIS rates are used to fit the cross-section of OIS rates. Moreover, adding additional OIS rates is not warranted given that they are included to improve the fit of model-implied interest rate expectations and that longer-maturity OIS rates contain significant term premia (see chapter 2).

Pricing Factors

Of additional interest for the OIS-augmented model is whether the inclusion of OIS rates affects the model's pricing factors \mathbf{x}_t . To investigate this, I compare the estimated principal components of the bond yields — used as pricing factors in the OLS/ML model — to the estimated model-implied pricing factors from Kalman filter estimation of the OIS-augmented models. Figure 3.4 plots the time series of the first three principal components, estimated from the panel of

Figure 3.4: Estimated Principal Components of the Actual Bond Yields and Estimated Pricing Factors from the 4-OIS-Augmented Model



Note: Estimated principal components from the actual bond yield data with the following maturities: 3, 6, 12, 18, 24, 30, 36, 42, 48, 54, 60, 84 and 120 months. Estimated pricing factors from the three-factor 4-OIS-augmented model, implied by the Kalman filter.

bond yields, and the estimated pricing factors from the 4-OIS-augmented model. For all three factors, the Kalman filter-implied pricing factors are nearly identical to the estimated principal components.³⁰ This implies that OIS rates do not include any additional information, over and above that in bond yields, of value in fitting the actual yields. This, again, is intuitive: OIS rates are included in the GADTSM to provide information useful for the identification of the risk-neutral yields, not the fitted yields.

Parameter Estimates

Recall, from section 3.3.2, that informational insufficiency in GADTSMs gives rise to finite sample bias. Persistent yields will have persistent pricing factors, resulting in estimates of the persistence parameters $\hat{\Phi}$ that are biased downwards. Following Bauer et al. (2012), I numerically assess the extent to which OIS-augmentation reduces finite sample bias by reporting the maximum eigenvalues of the estimated persistence parameters $\hat{\Phi}$. The higher the maximum eigenvalue, the more persistent the estimated process.

³⁰Table B.3, in appendix B.4.1, demonstrates that the summary statistics of the estimated principal components and pricing factors are very similar too.

As a benchmark, the maximum absolute eigenvalue of $\hat{\Phi}$ for the unaugmented OLS/ML model is 0.9985. The maximum absolute eigenvalue of $\hat{\Phi}$ for the survey-augmented model is 0.9983. However, for the 4-OIS-augmented model, the corresponding figure is 0.9988, indicating that, in comparison to the unaugmented model, augmentation with OIS rates does serve to mitigate finite sample bias.³¹ This indicates that OIS-augmentation does help to resolve the informational insufficiency in GADTSMs, and its associated symptoms. However, to assess this more thoroughly, a comparison of model-implied interest rate expectations is necessary. A well-identified model should accurately reflect the evolution of interest rate expectations.

Appendix B.4.1 presents further evidence to indicate that the OIS-augmentation resolves the identification problem, by studying the stability of interest rate expectation estimates using different sample periods of data. Estimates of interest rate expectations (on a given date) from the unaugmented OLS/ML model are more liable to vary as the sample period changes. In contrast, estimates from the 4-OIS-augmented model are remarkably stable, indicating that OIS-augmentation does provide useful information for the identification of interest rate expectations, which improves the model’s usefulness for real-time policy analysis.

3.6.2 Model-Implied Interest Rate Expectations

The central focus of this chapter is the identification and estimation of interest rate expectations within GADTSMs. Panels A and B of figure 3.5 plot the 2-year risk-neutral yields and term premia from the GADTSMs estimated between January 2002 and December 2016, respectively.

Panel A offers illustrative evidence of the effect of OIS-augmentation on the GADTSM estimates of expected future short-term interest rates. Over the 2002-2016 sample, the five models exhibit similar qualitative patterns, rising to peaks and falling to troughs at similar times. However, there are a number of notable differences between the series that help to illustrate the benefits of OIS-augmentation.

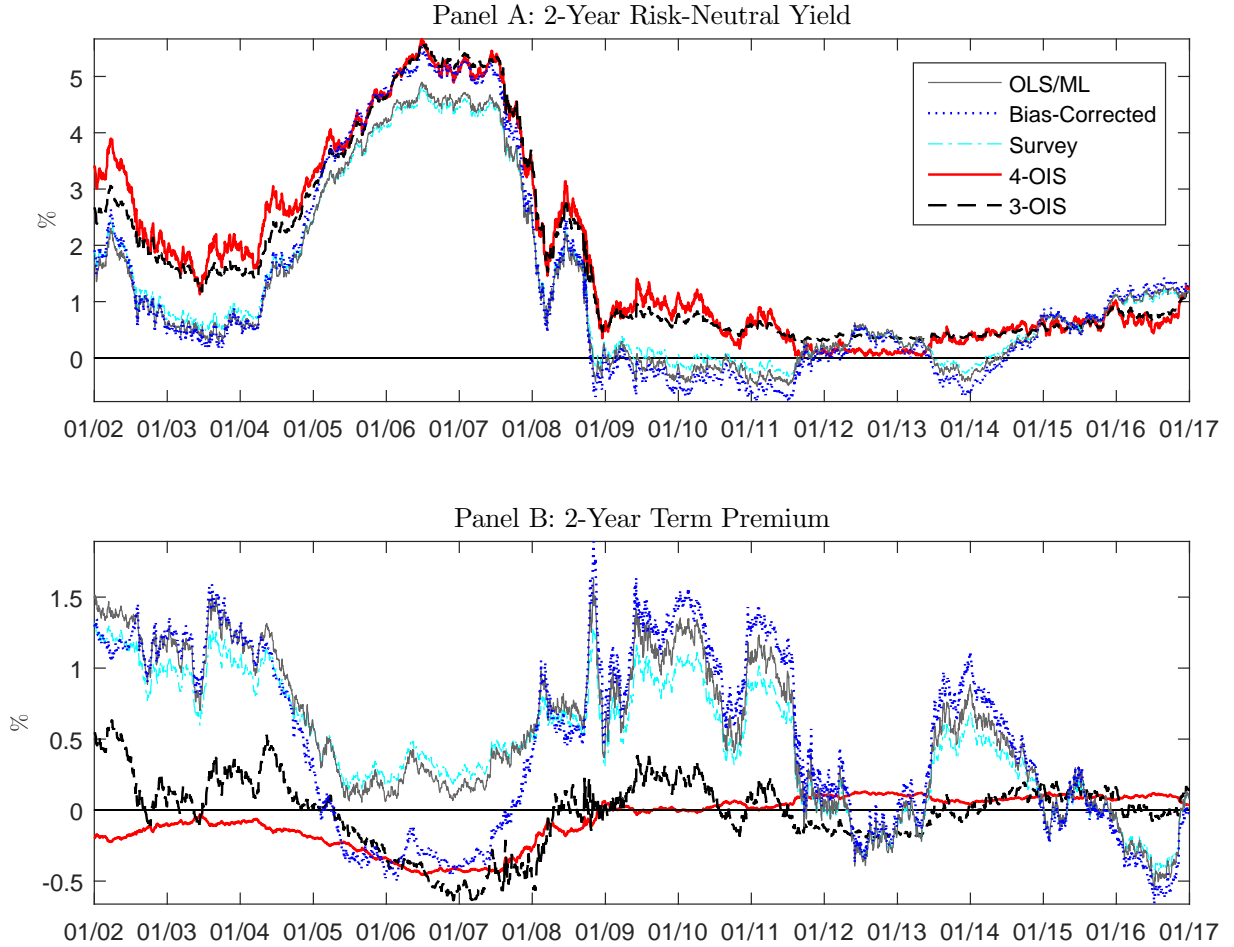
For the majority of the 2002-2016 sample, the OIS-augmented models generate 2-year risk-neutral yields that exceed those from the OLS/ML and bias-corrected models.³² Moreover, marked differences exist in the evolution of risk-neutral yield estimates from the models from late-2008 onwards. These differences have counterfactual implications for the efficacy of monetary policy at this time. First, from late-2008 to late-2011, the risk-neutral yields from the OLS/ML and bias-corrected models are persistently negative, implying counterfactual expectations of negative interest rates. In contrast, unlike the other models, the risk-neutral yields implied by the 3 and 4-OIS-augmented models obey a zero lower bound, with estimated interest rate expectations never falling negative, despite the fact that additional restrictions are not imposed to achieve this. This is true at all horizons, and represents an important contribution in the light of recent computationally burdensome proposals for term structure modelling at the zero lower bound (e.g. Christensen and Rudebusch, 2013a,b).

Second, between mid-2011 and 2013, the 2-year risk neutral yields from the OLS/ML and bias-corrected models reach a peak, indicating an increase in expected future short-term interest

³¹The corresponding statistic for the bias-corrected model, which performs bias-correction directly on the estimated Φ , is 1.0000 (to four decimal places). However, the ‘true’ pricing factor persistence is unknown.

³²Longer-horizon (i.e. 10-year) risk-neutral yields from the OIS-augmented models also exceed those from the OLS/ML and bias-corrected models. This is consistent with Meldrum and Roberts-Sklar (2015), who argue that unaugmented models provide “implausibly low estimates of long-term expected future short-term interest rates” (p. 1), “which in turn means that long-maturity term premium estimates are likely to be too high” (p. 3).

Figure 3.5: Estimated Yield Curve Decomposition



Note: Estimated risk-neutral yields (panel A) and term premia (panel B) from each of five GADTSMs, respectively. The five models are: (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML); (ii) the bias-corrected model (Bias-Corrected); (iii) the survey-augmented model (Survey); (iv) the 4-OIS-augmented model (4-OIS); and (v) the 3-OIS-augmented model (3-OIS). The models are estimated with three pricing factors, using daily data from January 2002 to December 2016. All figures are in annualised percentage points.

rates over a 2-year horizon. In contrast, during the same period, the 2-year risk-neutral yield estimates from the 3-OIS-augmented model remain broadly stable, while the corresponding estimates from the 4-OIS-augmented model fall slightly to a trough. From August 2011, the Federal Reserve engaged in calendar-based forward guidance designed to influence investors' expectations of future short-term interest rates, signalling that interest rates would be kept at a low level for an extended period of time. For instance, on August 9, 2011, the Federal Open Market Committee (FOMC) stated that it expected to keep the federal funds rate near zero "at least through mid-2013." This, and other forward guidance statements, were effective at deferring investors' expectations of future rate rises between mid-2011 and 2013. Swanson and Williams (2014) show that private sector expectations of the time until a US rate rise, from Blue Chip surveys, jumped from between 2 and 5 quarters to 7 or more quarters. In this respect, the finding that expectations of future short-term interest rates increased during this period, as implied by the OLS/ML and bias-corrected models, wrongly suggests that forward guidance policy was counterproductive. These models predict that investors began to expect rate rises sooner rather than later. The OIS-augmented models imply that investors were expecting rate rises no sooner, and possibly slightly later, than they had in previous period. Subsequent

quantitative analysis further demonstrates that the OIS-augmented models provide superior estimates of interest rate expectations during this period.

In figure 3.5, the estimated 2-year term premium from the 4-OIS-augmented model is persistently negative from 2002 to 2008. This is a direct consequence of the accurate fitting of risk-neutral yields. However, this feature is not true for all maturities; the estimated term premia at longer-horizons are frequently and persistently positive. For instance, the 10-year term premium from the 4-OIS-augmented model peaks at 79 basis points in late-2008.

Risk-Neutral Yields and Federal Funds Futures

To quantitatively evaluate the GADTSM-implied risk-neutral yields, I first compare them to expectations implied by 1 to 11-month federal funds futures contracts with matching horizon. Chapter 2 discusses how federal funds futures have long been used as measures of investors' expectations of future short-term interest rates. To facilitate the comparison, I first calculate 1, 2, ..., 11-month risk-neutral yields using the estimated model parameters from each GADTSM. I then calculate risk-neutral 1-month *forward* yields using the estimated risk-neutral yields.³³ Like federal funds futures contracts, the risk-neutral 1-month forward rates settle based on outcomes during a 1-month period in the future. However, because of the settlement structure of federal funds futures contracts, I only compare risk-neutral forward yields and federal funds futures rates on the final day of each calendar month.³⁴ I find that the risk-neutral forward yields from the OIS-augmented models most closely align with the expectations implied by federal funds futures rates with matching horizon, implying that they provide superior estimates of investors' expected future short-term interest rates compared to other GADTSMs.

Table 3.1 provides formal evidence to support this conclusion, presenting the RMSE of risk-neutral 1-month forward yields from different GADTSMs and corresponding-horizon federal funds futures rates. On a RMSE basis, the OIS-augmented models unambiguously provide superior estimates of expected future short-term interest rates, as measured by federal funds futures rates, at every horizon. The 4-OIS-augmented model outperforms the 3-OIS-augmented model at all but two horizons. Even at extremely short horizons, the benefits of OIS-augmentation are large: the RMSE fit of the unaugmented OLS/ML and bias-corrected models at the 3 to 4 month horizon is over three times larger than that of the 4-OIS-augmented model.

Despite fitting federal funds futures-implied interest rate expectations worse than the OIS-augmented models, the survey-augmented model does perform better than the unaugmented OLS/ML and bias-corrected models in this regard. This supports the claim that, while survey-augmentation does help to reduce the informational insufficiency problem in GADTSMs, quarterly frequency survey expectations are not sufficient for the accurate identification of interest rate expectations at higher frequencies.

³³To calculate the risk-neutral forward rate \tilde{f}_{t_1, t_2} from day t_1 to day t_2 , I use the following formula:

$$(1 + \tilde{y}_2)^{d_2} = (1 + \tilde{y}_1)^{d_1} (1 + \tilde{f}_{t_1, t_2})^{d_2 - d_1}$$

where \tilde{y}_1 (\tilde{y}_2) is the risk-neutral yield for the time period $(0, t_1)$ ($(0, t_2)$) and d_1 (d_2) is the length of time between time 0 and time t_1 (t_2) in years.

³⁴See chapter 2 for a detailed description of the settlement structure of federal funds futures contracts. The salient point is that an n -month federal funds futures contract traded on day t_j of the calendar month t has the same settlement period as an n -month contract traded on a different day t_k in the same calendar month t . For this reason, the horizon of a federal funds futures contract and the risk-neutral forward yield only align on the final calendar day of each month.

Table 3.1: GADTSM-Implied Expectations: Root Mean Square Error (RMSE) of the Risk-Neutral 1-Month Forward Yields *vis-à-vis* Corresponding-Horizon Federal Funds Futures Rates

Sample: January 2002 to December 2016					
Horizon	OLS/ML	BC	Survey	3-OIS	4-OIS
0 to 1 Months	0.2230	0.2196	0.2025	0.1649	0.1823
1 to 2 Months	0.2134	0.2049	0.1802	0.1201	0.1293
2 to 3 Months	0.2307	0.2167	0.1842	0.1205	0.0929
3 to 4 Months	0.2678	0.2496	0.2147	0.1394	0.0828
4 to 5 Months	0.3099	0.2888	0.2534	0.1552	0.0898
5 to 6 Months	0.3573	0.3344	0.2981	0.1593	0.1021
6 to 7 Months	0.4103	0.3865	0.3480	0.1576	0.1176
7 to 8 Months	0.4648	0.4422	0.3983	0.1545	0.1316
8 to 9 Months	0.5243	0.5030	0.4528	0.1530	0.1429
9 to 10 Months	0.9773	0.9599	0.9200	0.6755	0.6564
10 to 11 Months	1.3392	1.3297	1.2776	0.9915	0.9635

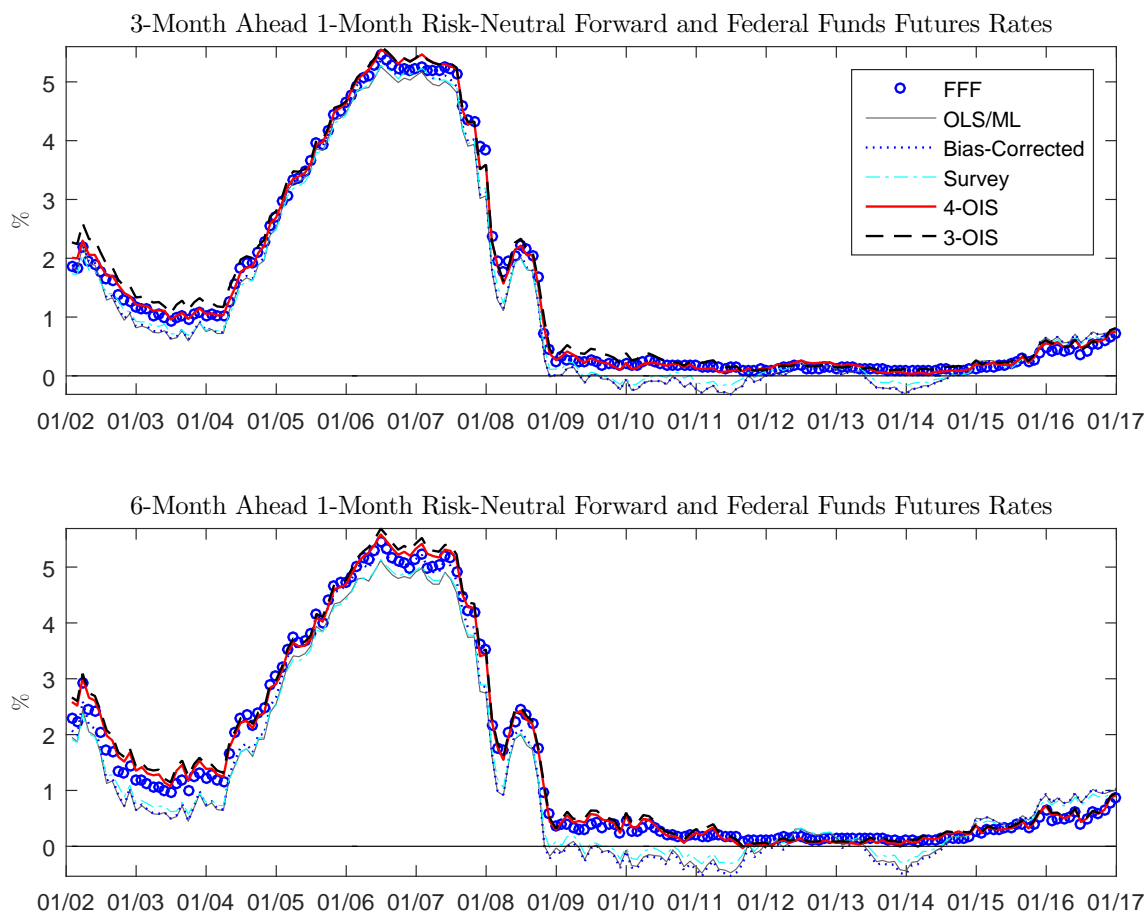
Note: RMSE of the risk-neutral 1-month forward yields from each of the five GADTSMs in comparison to corresponding-horizon federal funds futures rates. The five models are: (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML); (ii) the bias-corrected model (Bias-Corrected); (iii) the survey-augmented model (Survey); (iv) the 3-OIS-augmented model (3-OIS); and (v) the 4-OIS-augmented model (4-OIS). The models are estimated with three pricing factors, using daily data from January 2002 to December 2016. The risk-neutral forward yields and the federal funds futures rates are compared on the final day of each calendar month. All figures are in annualised percentage points. The lowest RMSE model at each maturity has been emboldened for ease of reading.

Figure 3.6 provides visual comparison of the risk-neutral 1-month forward yields and federal funds futures rates at the 3 to 4 and 6 to 7 month horizons. Here, I plot the risk-neutral 1-month forward yields from the unaugmented OLS/ML, bias-corrected, survey-augmented, 4-OIS-augmented and 3-OIS-augmented GADTSMs, as well as corresponding-horizon federal funds futures rates. The plot highlights the causes of the differences in fit highlighted by table 3.1. Two important observations follow.

First, in the pre-crisis period (to mid-2007), the OLS/ML and survey-augmented GADTSMs generate estimated risk-neutral forward yields that persistently fall below the corresponding-horizon federal funds futures rates. The bias-corrected risk-neutral forward yields also fall below the corresponding-horizon federal funds futures rates until 2005. In contrast, the estimated risk-neutral forward yields from the OIS-augmented models align more closely with federal funds futures rates during this period, especially at the 3 to 4-month horizon.

Second, from late-2008 until late-2011, and from mid-2013 to mid-2014, the risk-neutral forward yields from the OLS/ML, bias-corrected, and survey-augmented models differ greatly from the corresponding-horizon federal funds futures rates. Moreover, these models offer counterfactual predictions for the evolution of interest rate expectations during this period. In particular, in these periods, the risk-neutral forward yields from the OLS/ML, bias-corrected and survey-augmented models are persistently negative, implying that investors expected future short-term interest rates to fall negative. Moreover, from late-2011 to mid-2012, the risk-neutral forward yields from the OLS/ML and bias-corrected models rise to a peak. Not only is this contrary to the policy narrative at the time — policymakers were engaging in calendar-based forward guidance that sought to push back the date investors expected policy rates to lift-off from their lower bound — it is also counterfactual with respect to market-implied interest rate expecta-

Figure 3.6: Estimated Risk-Neutral 1-Month Forward Yields and Comparable-Horizon Federal Funds Futures (FFF) Rates



Note: Estimated 3 month (top panel) and 6 month (bottom panel) ahead 1-month risk-neutral forward yields from each of five GADTSMs. The five models are: (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML); (ii) the bias-corrected model (Bias-Corrected); (iii) the survey-augmented model (Survey); (iv) the 4-OIS-augmented model (4-OIS); and (v) the 3-OIS-augmented model (3-OIS). The models are estimated with three pricing factors, using daily data from January 2002 to December 2016. I compare the estimated risk-neutral forward yields to corresponding-horizon federal funds futures (FFF) rates, plotted on the final day of each calendar month. All figures are in annualised percentage points.

tions. In contrast, the OIS-augmented models align closely with federal funds futures-implied expectations for most of the 2009-2016 period.

Overall, the comparison of risk-neutral forward yields and federal funds rates supports the claim that OIS-augmentation improves the identification of interest rate expectations.

Risk-Neutral Yields and Short-Horizon Survey Expectations

As further evidence in support of this claim, I compare the model-implied interest rate expectations to short-horizon survey expectations. The preferred GADTSM(s) should also be able to reasonably capture the qualitative and quantitative evolution of comparable-horizon survey expectations. Against this metric, I find that the OIS-augmented model provides superior overall estimates of short-term interest rate expectations, in comparison to all other models.

I compare the estimated 1.5, 4.5, 7.5, 10.5 and 13.5-month risk-neutral yields to corresponding-horizon survey expectations. I calculate approximate short-term interest rate expectations using SPF data for the median expectation of the 3-month T-Bill rate for the remainder of the current

Table 3.2: GADTSM-Implied Expectations: Root Mean Square Error (RMSE) of the In-Sample Risk-Neutral Yields *vis-à-vis* 1.5, 4.5, 7.5, 10.5 and 13.5-Month Survey Expectations

Sample: January 2002 to December 2016					
RMSE vs. Survey Expectation at Different Horizons					
Model	1.5-Month	4.5-Month	7.5-Month	10.5-Month	13.5-Month
OLS/ML	0.1853	0.2005	0.2700	0.3639	0.4744
Bias-Corrected	0.1857	0.1979	0.2643	0.3570	0.4686
Survey	0.1749	0.1661	0.2216	0.3087	0.4135
3-OIS	0.1642	0.1437	0.1514	0.1545	0.1634
4-OIS	0.1727	0.1354	0.1199	0.1311	0.1577

Note: RMSE of the risk-neutral yields from each of the five GADTSMs in comparison to approximated *Survey of Professional Forecasters* survey expectations, using estimated risk-neutral yields on SPF deadline dates. The five models are: (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML); (ii) the bias-corrected model (Bias-Corrected); (iii) the survey-augmented model (Survey); (iv) the 3-OIS-augmented model (3-OIS); and (v) the 4-OIS-augmented model (4-OIS). The models are estimated with three pricing factors, using daily data from January 2002 to December 2016. The construction of the survey expectation approximations is described in appendix A.2. All figures are in annualised percentage points. The lowest RMSE model at each maturity has been emboldened for ease of reading.

quarter, and in the first, second, third and fourth quarters ahead. A complete description of how these expectations are approximated is presented in appendix A.2. To compare the estimated risk-neutral yields to these survey expectations, I calculate the RMSE of the risk-neutral yields *vis-à-vis* the corresponding-horizon survey expectation on survey submission deadline dates.³⁵

Table 3.2 shows that, on a RMSE basis, the OIS-augmented models unambiguously provide superior estimates of expected future short-term interest rates at each horizon. By this metric, the OLS/ML and bias-corrected models provide the most inferior estimates of future short-term interest rate at all three horizons.

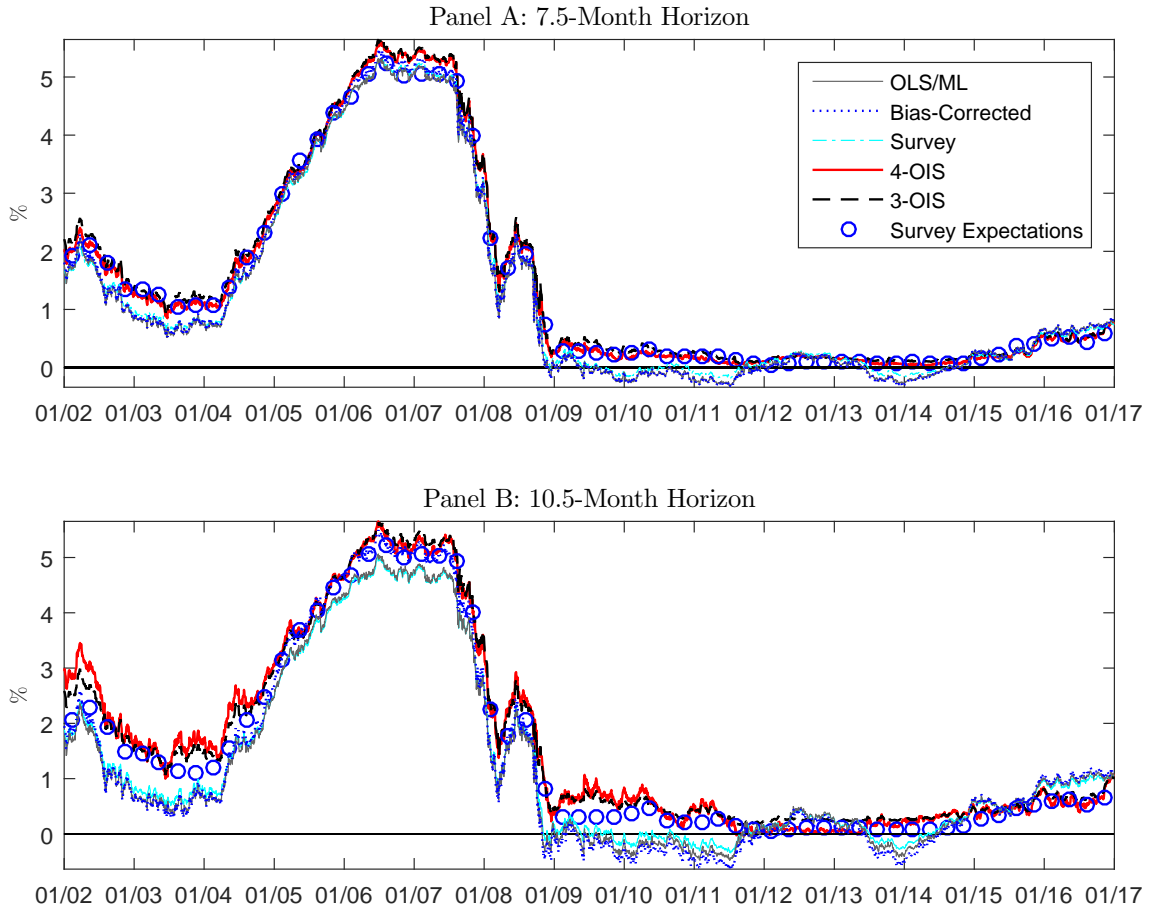
At the 4.5, 7.5, 10.5 and 13.5-month horizons the 4-OIS-augmented model provides the superior fit of survey expectations. Strikingly, at the 10.5 and 13.5-month horizons, the RMSE fit of the OLS/ML and bias-corrected models are around three times the RMSE fit of the 4-OIS-augmented model. Although the 3-OIS-augmented model provides the lowest RMSE fit for the 1.5-month survey expectation, the RMSE fit of the 4-OIS-augmented model is only 0.85 basis points higher at this horizon. In contrast, at the 10.5-month horizon the RMSE fit of the 4-OIS-augmented model is 2.34 basis points lower than the 3-OIS-augmented model.

Surprisingly, the survey-augmented model, which uses the same SPF survey expectations as an input to estimation, does not provide a superior fit for these expectations at any horizon *vis-à-vis* the OIS-augmented models. This supports the claim that quarterly frequency survey expectations are not sufficient for the accurate identification of higher frequency interest rate expectations within a GADTSM framework. Nevertheless, the RMSE fit of the survey-augmented model is superior to the fit of both the OLS/ML and bias-corrected models at all horizons, supporting the claim that augmentation of GADTSMs with additional information can aid the identification of risk-neutral yields.

Figure 3.7 graphically illustrates the evolution of estimated risk-neutral yields and the ap-

³⁵There are two caveats to this comparison which help to explain small differences between survey expectations and risk-neutral yields. First, the expectational horizons of risk-neutral yields and the T-Bill expectations do not exactly correspond, because the latter 3-month T-Bill rate expectations still reflect expected developments up to 3 months beyond the horizon. Second, 3-month T-Bill rates are on a discount basis, whereas risk-neutral yields are quoted on a yield basis.

Figure 3.7: Short-Term Interest Rate Expectations



Note: Estimated 7.5-month (panel A) and 10.5-month (panel B) risk-neutral yields from each of five GADTSMs. The five models are: (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML); (ii) the bias-corrected model (Bias-Corrected); (iii) the survey-augmented model (Survey); (iv) the 4-OIS-augmented model (4-OIS); and (v) the 3-OIS-augmented model (3-OIS). The models are estimated with three pricing factors, using daily data from January 2002 to December 2016. I compare the estimated risk-neutral yields to approximated survey expectations of future short-term interest rates over the same horizon. The construction of the survey expectation approximations, using data from the *Survey of Professional Forecasters*, is described in appendix A.2. All figures are in annualised percentage points.

proximated survey expectations at the 7.5 and 10.5-month horizons — the plots for the 1.5, 4.5 and 13.5-month horizons are qualitatively similar. Three observations follow.

First, between 2002 and late-2004, the OLS/ML, bias-corrected and survey-augmented models generate estimated risk-neutral yields that persistently fall below the corresponding-horizon survey expectation. In contrast, the estimated risk-neutral yields from the OIS-augmented models closely co-move with the approximated survey expectations during this period, especially at the 7.5-month horizon. This corroborates with the comparison of risk-neutral forward yields and federal funds future-implied interest rate expectations in section 3.6.2.

Second, between early-2006 and mid-2007, the risk-neutral yields from the OIS-augmented models exceed interest rate expectations implied by surveys. During this short period, the 7.5-month risk-neutral yields from the OLS/ML, bias-corrected and survey-augmented models more closely align with survey expectations. Although this is not true at the 10.5-month horizon, where the OIS-augmented and bias-corrected models perform best in this period. Recall from figure 3.6 that the risk-neutral forward yields from the OIS-augmented models closely align with federal funds futures-implied interest rate expectations during this period.

Third, as in section 3.6.2, the GADTSMs offer markedly different estimates of interest rate expectations from late-2008 to late-2011. During this period, the 7.5 and 10.5-month risk-neutral yields from the OLS/ML and bias-corrected models are persistently negative, implying, counter-factually, that investors expected future short-term interest rates to fall negative. From late-2011 to mid-2012, the risk-neutral yields from the OLS/ML and bias-corrected models rise to peak. Again, this is both contrary to the policy narrative and the survey expectations at the time. In contrast, the OIS-augmented models — the 4-OIS-augmented model especially — align closely with survey expectations for much of the post-2008 period.

Overall, the results further support the claim that the OIS-augmentation of GADTSMs improves the identification of interest rate expectations. At short-term horizons, OIS-augmented models provide superior estimates of investors' expectations of future short-term interest rates for much of the 2002-2016 sample.

Risk-Neutral Yields and Long-Horizon Survey Expectations

The expectational horizons considered in the previous sub-section are short-term. However, GADTSMs provide estimates of risk-neutral yields for the whole term structure, at horizons further into the future. This is an important motive for using GADTSMs to estimate interest rate expectations, instead of market-based financial measures; market-based financial measures seldom provide accurate measures of investors' interest rate expectations at horizons in excess of 2 years (see chapter 2).

Within a GADTSM, the 10-year risk-neutral yield on date t provides an estimate for the expected average short-term interest rate for the 10-year period following date t . In general, survey data on these longer-term interest rate expectations are not readily available, making it difficult to systematically test the long-horizon interest rate expectations attained from GADTSMs. However, in recent years, the New York Federal Reserve's *Survey of Primary Dealers* has asked respondents an increasing number of questions regarding their longer-horizon interest rate expectations.³⁶ Specifically, since October 2013, respondents have been asked to: “provide your estimate of the longer run target federal funds rate and your expectation for the average federal funds rate over the next 10 years”.³⁷ The latter of these requests corresponds to the information contained within the 10-year risk-neutral yields attained from the GADTSMs: the expectation of the average of the short-term interest rate over a 10-year horizon.

To quantitatively assess the longer-horizon interest rate expectations implied by the GADTSMs, I compare the estimated 10-year risk-neutral yield to the median “expectations for the average federal funds rate over the next 10 years” of survey respondents on the survey deadline dates. Again, I calculate the RMSE fit of the risk-neutral yields *vis-à-vis* the survey expectations.

Table 3.3 presents the results from this analysis. Although the sample of long-horizon survey expectations is relatively short, including 26 surveys from October 2013 to December 2016, the results support the primary conclusion of this chapter: that the OIS-augmented models provide unambiguously superior estimates of future short-term interest rate expectations. The RMSE fit of the OLS/ML, bias-corrected and survey-augmented models are over double the RMSE fit of the 4-OIS-augmented model. Moreover, the RMSE fit of the 4-OIS-augmented is smaller

³⁶The questions and results of these surveys are publicly available from: www.newyorkfed.org/markets/primarydealer_survey_questions.html.

³⁷In the surveys, the question preceding this was: “provide your estimate of the most likely outcome (i.e., the mode) for the target federal funds rate or range at the end of each half-year period”.

Table 3.3: GADTSM-Implied Expectations: Root Mean Square Error (RMSE) of the In-Sample Risk-Neutral Yields *vis-à-vis* 10-Year Survey Expectation

Model	Sample: October 2013 to December 2016	
	RMSE vs. 10-Year Expectation, Survey of Primary Dealers	
OLS/ML	1.5878	
Bias-Corrected	1.8747	
Survey	1.5309	
3-OIS	0.7831	
4-OIS	0.7034	

Note: RMSE of the risk-neutral yields from each of the five GADTSMs in comparison to the 10-year survey expectation, using estimated risk-neutral yields on survey deadline dates. The five models are: (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML); (ii) the bias-corrected model (Bias-Corrected); (iii) the survey-augmented model (Survey); (iv) the 3-OIS-augmented model (3-OIS); and (v) the 4-OIS-augmented model (4-OIS). The models are estimated with three pricing factors, using daily data from January 2002 to December 2016. The median survey expectation is from the Survey of Primary Dealers, New York Federal Reserve. All figures are in annualised percentage points. The lowest RMSE model at each maturity has been emboldened for ease of reading.

than that of the 3-OIS-augmented model, indicating that longer-horizon OIS rates do help to improve the model's fit of interest rate expectations at longer tenors.

Daily Changes in GADTSM-Implied Interest Rate Expectations

OIS-augmentation offers benefits for the identification and estimation of interest rate expectations from GADTSMs at a *daily frequency*. As OIS rates are available at a daily frequency, they offer potentially sizeable benefits when estimating the daily frequency evolution of interest rate expectations. To illustrate these benefits, I directly analyse the daily changes in GADTSM-implied risk-neutral yields.

The analysis of daily changes in interest rate expectations is an integral part of historical monetary policy analysis. Most recently, a number of authors have used daily changes in interest rate expectations and term premia to assess the relative efficacy of various interest rate channels of unconventional monetary policies (see chapter 4 and the references within). For the OIS-augmented GADTSM to be well-suited to historical policy analysis of this sort, it is important that the risk-neutral yields provide an accurate depiction of the daily frequency evolution of interest rate expectations. Specifically, for the GADTSM-implied interest rate expectations to reasonably reflect the expectations of investors over a comparable horizon at a daily frequency, they should, at the very least, qualitatively match numerical measures of investors' interest rate expectations. To test this, I compare the sign of daily changes in 3, 6 12 and 24-month risk-neutral yields to the sign of daily changes in comparable-maturity OIS rates.³⁸ For the GADTSM to reasonably reflect investors' expectations, the sign of the daily change in the risk-neutral yield should correspond to the sign of the daily change in the comparable horizon OIS rate. I record the proportion of positive and negative daily changes in OIS rates that are matched in sign by the change in the corresponding-horizon risk-neutral yields. To focus on significant

³⁸I use the sign of daily changes in OIS rates because their horizon corresponds exactly to that of the nominal government bond yields I use. Although it may seem somewhat tautological to compare an OIS-augmented GADTSM to OIS rates, previous results indicate that this need not be the case. In table 3.2, the survey-augmented model does not provide the best fit for the survey-expectations which are used as an input to its estimation.

Table 3.4: Proportion of Daily Changes in OIS Rates Matched in Sign by the Daily Changes in In-Sample GADTSM Risk-Neutral Yields

Sample: January 2002 to December 2016				
Model	Maturity			
	3-Months	6-Months	1-Year	2-Years
Proportion of Positive Daily Changes Matched				
OLS/ML	84.33%	93.20%	95.10%	94.27%
Bias-Corrected	83.58%	92.72%	95.10%	95.13%
Survey	82.09%	91.26%	92.81%	95.13%
3-OIS	85.82%	92.72%	94.44%	96.28%
4-OIS	85.82%	94.17%	96.08%	97.99%
Proportion of Negative Daily Changes Matched				
OLS/ML	87.21%	93.20%	94.93%	93.18%
Bias-Corrected	86.63%	93.13%	95.22%	93.18%
Survey	86.05%	91.85%	94.93%	96.85%
3-OIS	90.12%	93.99%	94.93%	98.69%
4-OIS	90.70%	93.99%	96.42%	98.69%

Note: Proportion of daily changes in 3, 6, 12 and 24-month OIS rates (in excess of one standard deviation of their daily change in absolute value) that are matched in sign by the daily change in the corresponding maturity GADTSM risk-neutral yield. All proportions are expressed as a percentage to two decimal places. Five GADTSMs are compared: (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML); (ii) the bias-corrected model (Bias-Corrected); (iii) the survey-augmented model (Survey); (iv) the 3-OIS-augmented model (3-OIS); and (v) the 4-OIS-augmented model (4-OIS). The models are estimated with three pricing factors, using daily data from January 2002 to December 2016. The highest percentage model at each maturity has been emboldened for ease of reading.

changes in OIS rates, I omit days on which OIS rates changed by less, in absolute value, than one standard deviation of the daily changes in the OIS rate over the whole sample (approx 2-5 basis points). The results are presented in table 3.4.

The results indicate that the 4-OIS-augmented model provides the best qualitative match for the sign of daily changes in 3, 6, 12 and 24-month OIS rates. For example, the 4-OIS-augmented model is the only to match over 96% of positive daily changes in 1-year OIS rates. Moreover, at the 2-year horizon, the sign of daily changes in the risk-neutral yield from the 4-OIS-augmented model matches 97.99% (98.69%) of positive (negative) OIS rate, around 3 (5) percentage points more than the OLS/ML and bias-corrected models match.

Overall, the results in table 3.4 are consistent with the claim that the 4-OIS-augmented model best reflects the daily frequency evolution of short-term interest rate expectations.

3.6.3 Explaining the Benefits of OIS-Augmentation

The preceding discussion highlights that OIS-augmented models provide estimates of expected future short-term interest rates that are superior to the OLS/ML, bias-corrected and survey-augmented models. Moreover, within the class of OIS-augmented models considered, the 4-OIS-augmented model, on balance, outperforms the 2 and 3-OIS-augmented models.

Figure 3.7 highlights that there are differences between the risk-neutral yields from OIS-augmented models and the OLS/ML and bias-corrected models for the whole 2002-2016 sample period. In particular, since late-2008, the risk-neutral yields from the models offer distinctly different qualitative and quantitative predictions for estimated interest rate expectations.

To understand the economic reasons behind these differences, I draw on the canonical de-

scription of the first three principal components of bond yields as the level, slope and curvature of the yield curve respectively, together with the model-implied loadings on these factors.³⁹ Figure 3.8 plots these loadings for both the calculation of fitted yields $B_n \equiv -\frac{1}{n}\mathcal{B}_n(\boldsymbol{\delta}_1, \boldsymbol{\Phi}^Q; \mathcal{B}_{n-1})$ (top row) and the risk-neutral yields $\tilde{B}_n \equiv -\frac{1}{n}\mathcal{B}_n(\boldsymbol{\delta}_1, \boldsymbol{\Phi}; \mathcal{B}_{n-1})$ (bottom row) for the 3-month to 10-year maturities. To refine discussion, loadings are presented for the two most inferior models — OLS/ML and bias-corrected — and the most superior — 4-OIS-augmented models. These loadings illustrate the extent to which the fitted and risk-neutral yields react to a one unit shock to a pricing factor at a given maturity, keeping all other pricing factors constant.

Unsurprisingly, the loadings for the fitted yields from the OLS/ML, bias-corrected and 4-OIS-augmented models are almost identical at all maturities, reinforcing the similarities in their fitted yields. The benefits of OIS-augmentation arise from the separate identification of interest rate expectations and term premia, rather than the fitting of actual yields.

However, the loadings for the risk-neutral yields differ at all horizons, helping to explain why the 4-OIS-augmented model is superior as a measure of interest rate expectations, and providing economic reasons for the differences in risk-neutral yields from late-2008 onwards. In particular, the risk-neutral loadings on the slope and curvature factors from the OIS-augmented model are close to zero for most maturities. In the OIS-augmented model, shocks to the level of the yield curve, such as the reduction of interest rates to their ELB in 2008, have the strongest effect on risk-neutral yields, helping to explain why OIS-augmented risk-neutral yields remain positive during the ELB period.

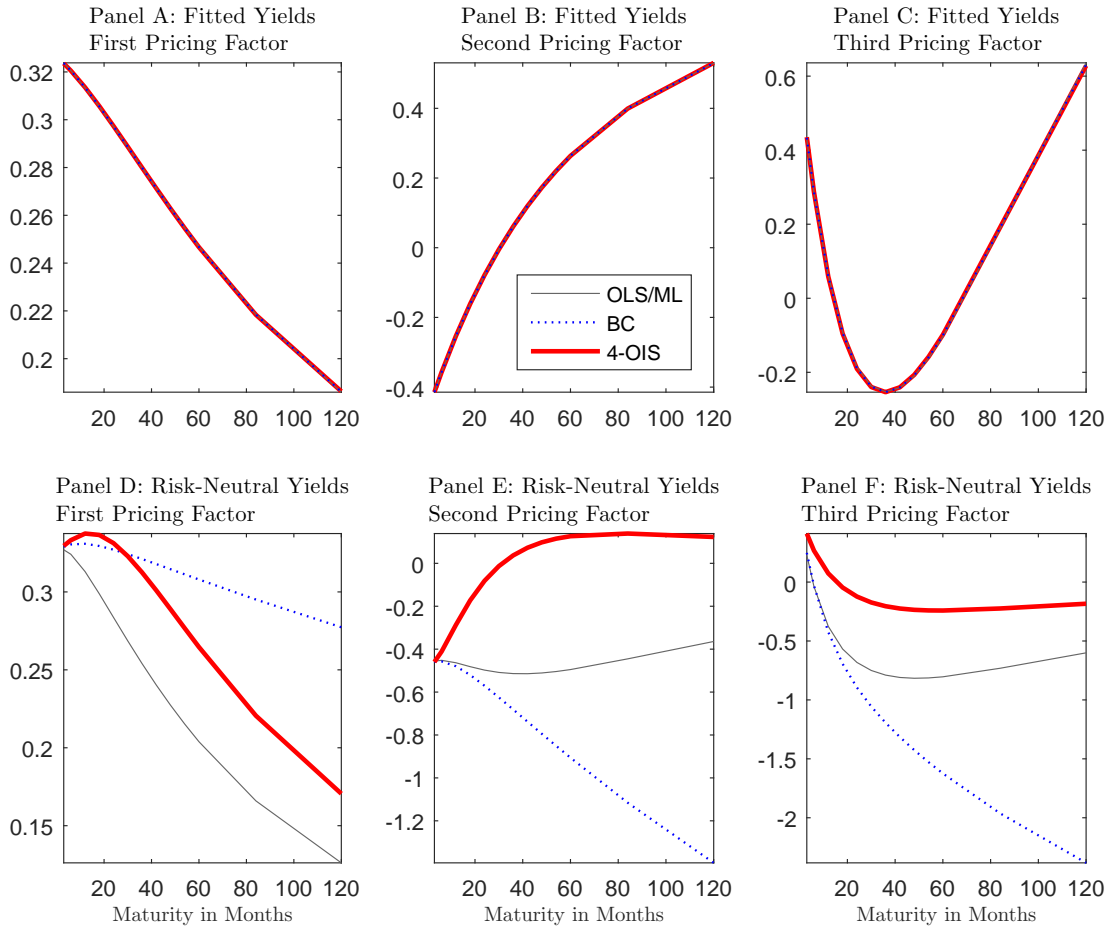
From late-2011 to 2013, the risk-neutral yields from the OLS/ML and bias-corrected models rise to a peak during 2012, falling back below zero for a short period from mid-2013 to mid-2014. The risk-neutral yields from the OIS-augmented models do not peak during 2012. This period was characterised by two notable phenomena. First, the target federal funds rate was at its effective lower bound. Having been set at this level in December 2008, the FOMC were signalling, through forward guidance, that it would be kept at this rate into the future. Second, the Eurozone sovereign debt crisis elevated Eurozone government bond yields. This was associated with a reduction in yields on, comparatively safe, longer-term US government bonds.⁴⁰ During the 2011-2013 period therefore, the US yield curve was characterised by a reduction in its *slope*, with no change in the *level* of short-term interest rates.

Panel E of figure 3.8 illustrates that a decrease in the slope of the yield curve places upward pressure on estimated risk-neutral yields in the OLS/ML and bias-corrected models at all maturities. That is, decreases in the yield curve slope, for a given level and curvature, tend to be associated with diminished term premia. However, the risk-neutral yields from the 4-OIS-augmented model react less strongly to a change in the yield curve slope, and, at longer-term horizons, a decrease in the yield curve slope will place downward pressure on estimated risk-neutral yields. This helps to explain why the risk-neutral yields from the 4-OIS-augmented do not rise to a peak in mid-2012, while those from the OLS/ML and bias-correct models do. The 4-OIS-augmented model does not exhibit the same peak, because the inclusion of OIS rates in the estimation alters the loading on that pricing factor. This constellation of factor loadings helps to attain risk-neutral yields from the 4-OIS-augmented model that align more closely with

³⁹Because the estimated pricing factors from the three-factor OIS-augmented models almost exactly correspond with the estimated principal components (see figure 3.4), this economic intuition is valid for these models.

⁴⁰Longer-term Eurozone interest rates peaked in 2011-2012 as a result of the sovereign debt crisis (Corsetti, Kuester, Meier, and Müller, 2013, 2014).

Figure 3.8: Fitted and Risk-Neutral Yield Factor Loadings



Note: I plot the estimated yield loadings B_n for the fitted and risk-neutral yields, for each the three pricing factors (level, slope and curvature respectively), from the OLS/ML, bias corrected and 4-OIS-augmented models estimated with three factors from January 2002 to December 2016. These coefficients can be interpreted as the *ceteris paribus* response of the fitted and risk-neutral bond yields at a given maturity to a contemporaneous shock to the respective pricing factor. The horizontal axis labels denote the maturity, in months. The three models are denoted by: (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML); (ii) the bias-corrected model (BC); and (iii) the 4-OIS-augmented model (4-OIS).

survey and market-implied expectations of future short-term interest rates.

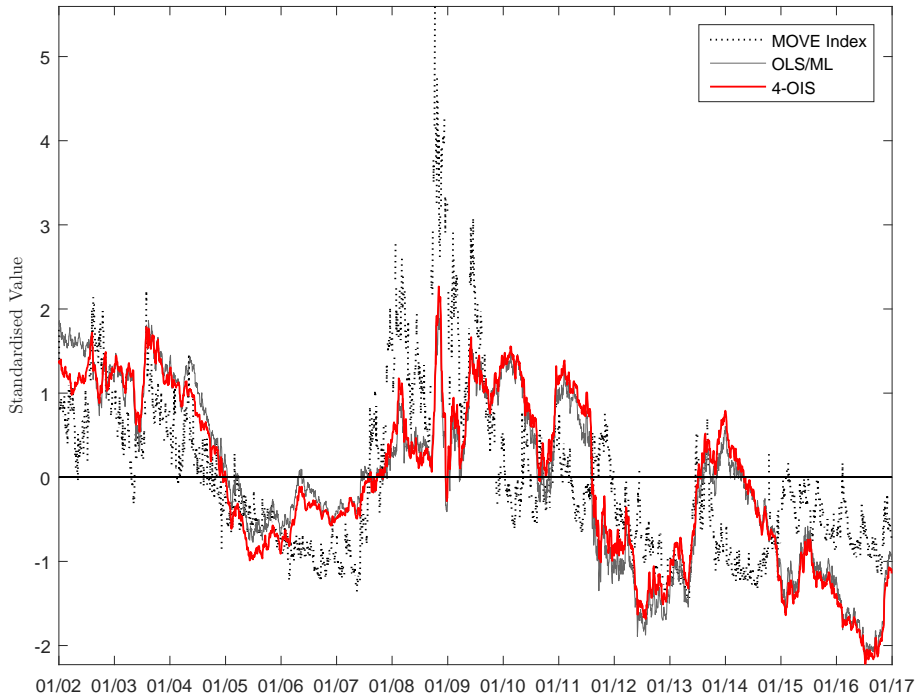
3.6.4 Model-Implied Term Premia

Alongside estimates of expectations of future short-term interest rates, GADTSMs provide estimates for the daily evolution of term premia. Although there is no direct metric against which to compare estimated term premia, Adrian et al. (2013) compare a standardised version of their estimated daily 10-year term premium to a standardised version of the 1-month Merrill Lynch Option Volatility Estimate (MOVE) index. This latter index is a measure of implied volatility from option contracts written on US Treasuring bonds.⁴¹ Thus, variation in MOVE reflects changes in the risk of holding US Treasuries.

Like Adrian et al. (2013), in figure 3.9 I plot the standardised z -score estimates of the 10-year term premium from the OLS/ML and 4-OIS-augmented models against the standardised

⁴¹Formally, the series used here (and in Adrian et al., 2013) is defined as a yield curve weighted index of the normalised implied volatility on 1-month Treasury options. It is the weighted average of volatilities on 2, 5, 10 and 30-year bond yields.

Figure 3.9: Standardised 10-Year Term Premia and Merrill Lynch Option Volatility Estimate (MOVE) Index



Note: I plot the standardised one-month Merrill Lynch Option Volatility Estimate (MOVE) index against the standardised estimates of term premia from (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML), and (ii) the 4-OIS-augmented model (4-OIS). The models are estimated with three factors from January 2002 to December 2016.

one-month MOVE index. The time series exhibit a strong positive correlation. The correlation coefficient between the standardised 10-year term premium estimate from the 4-OIS-augmented model and the standardised MOVE index is 0.62, marginally higher than the corresponding statistic of 0.60 for the OLS/ML model. This indicates that the estimated term premia from the 4-OIS-augmented model do reflect the risk of holding Treasury bonds.

3.7 Conclusion

Financial market participants and policymakers closely monitor the evolution of interest rate expectations using a wide range of financial market instruments. In this chapter, I investigate the informational content in OIS rates for this purpose and document how OIS rates can be used to improve estimates of nominal interest rate expectations, and term premia, attained from the term structure of nominal government bond yields.

Drawing on results in chapter 2, which demonstrate that 1 to 24-month US OIS rates provide accurate measure of investors' interest rate expectations, I present an OIS-augmented GADTSM for estimating the daily frequency evolution of interest rate expectations that explicitly accounts for the payoff structure in OIS contracts. Most existing arbitrage-free GADTSMs use information in the term structure of nominal government bond yields to identify both expectations of the future path of short-term interest rates and term premia. Numerous authors have drawn attention to an informational insufficiency in the estimation of these models. Kim and Orphanides (2012) propose survey-augmented GADTSMs as a solution to this problem. However, survey forecasts of future interest rates are only available at a low frequency (quarterly or monthly,

at best) and reflect investors' expectations of future short-term interest rates for a window of time in the future (e.g. one, two, three or four quarters ahead). OIS rates, on the other hand, are available at a daily frequency and have a horizon that aligns exactly with those of the zero-coupon nominal government bonds used in the estimation of GADTSMs. The term structure of OIS rates can therefore be readily added to a GADTSM for nominal government bond yields. I show that augmenting the GADTSM with OIS rates provides additional information, specifically related to future interest rate expectations, that can help better identify the evolution of these expectations. Using 3 to 24-month OIS rates in an arbitrage-free GADTSM enables the estimation of future short-term interest rate expectations for the whole term structure — from 3 months to 10 years. Estimates of interest rate expectations from OIS-augmented GADTSMs are superior to those from existing GADTSMs. In particular, short and long-horizon in-sample OIS-augmented risk-neutral yields match patterns in federal funds futures rates and survey expectations. These time series also match qualitative daily patterns exhibited by financial market instruments. This implies that OIS-augmented GADTSMs are well suited for daily frequency policy analysis. Thus, OIS-augmented GADTSMs provide reliable and policy-relevant estimates of interest rate expectations along the whole term structure.

Additionally, this chapter highlights the need to test the performance of GADTSMs in a range of dimensions — for example accuracy of fitted yields, risk-neutral yields and term premia — before applying them to analysis of monetary policy. This chapter proposes a battery of such tests for future research.

The contribution of this chapter extends beyond the GADTSM-literature. For example, the OIS-augmented GADTSM can be applied to better understand the transmission of monetary policy domestically and internationally, as chapters 4 and 5 demonstrate.

To conclude, OIS rates accurately reflect investors' near-term expectations of future short-term interest rates, providing useful information for improved identification of interest rate expectations and term premia at a range of horizons in arbitrage-free GADTSMs.

Chapter 4

Unconventional Monetary Policy and the Interest Rate Channel: Signalling and Portfolio Rebalancing

4.1 Introduction

Before the recent crisis, monetary policy was primarily conducted with one instrument: a short-term nominal interest rate. In the wake of financial turmoil and the subsequent reduction of short-term interest rates to their effective lower bound (ELB), central banks increasingly turned to ‘unconventional’ monetary policy tools, defined here as instruments beyond the ‘traditional’ policy rate. In this chapter, I focus on US unconventional monetary policies announced since November 2008:¹ large-scale asset purchases (LSAPs) and forward guidance. These two policies can transmit to the real economy, *inter alia*, via an interest rate channel with two sub-components: signalling and portfolio rebalancing. I assess the relative importance of these two channels for US unconventional policy in terms of their effect on the *real economy* — the ultimate goal of the policies. I show that unconventional monetary policies have placed significant downward pressure on long-term interest rates via both the signalling and portfolio balance channels. The primary finding is that reductions in long-term interest rates during the period of unconventional monetary policy easing between November 2008 and April 2013 have exerted a more powerful influence on the real economy through the signalling channel than through portfolio rebalancing. In terms of long-term bond yield and industrial production effects, the signalling channel is associated with around two-thirds to three-quarters of the total effects attributed to the two channels.

Federal Reserve (Fed) LSAPs have involved the direct purchase of longer-term assets from secondary markets. Since December 2008, the Fed has purchased a range of longer-term US Treasuries, agency debt and mortgage-backed securities (MBS), expanding its balance sheet by over 600%.² The policy was initiated “to put downward pressure on yields of a wide range of longer-term securities, support mortgage markets and promote a stronger economic recovery”.³

¹US large-scale asset purchases were first announced on November 25, 2008, just before the Federal Funds rate was lowered to its ELB on December 16, 2008.

²A similar policy was adopted by the Bank of England in March 2009, stimulating a large body of research in itself. The BoE has predominantly purchased longer-term UK gilts. The focus of this chapter, and the references within, is on US policy.

³www.federalreserve.gov/faqs/what-are-the-federal-reserves-large-scale-asset-purchases.htm.

The Fed announced the purchase of MBS and agency-backed bonds from private markets on November 25, 2008. On March 18, 2009 this was extended to include the purchase of \$300 billion of longer-term Treasury securities over a six-month period. These combined purchases were dubbed ‘QE1’ and concluded on March 16, 2010, with the Fed holding \$1.25 trillion of MBS and \$175 billion of agency-backed debt. The value of the asset stock was held constant until the inception of ‘QE2’ on November 3, 2010, following strong suggestions of further purchases in Fed Chairman Ben Bernanke’s August speech at Jackson Hole⁴ and his October speech at the Boston Fed.⁵ From the outset of QE2, the Fed stated that it would purchase \$600 billion of longer-term US Treasuries over a six-month period, concluding in June 2011. ‘QE3’ marked the most recent expansion of US LSAPs, announced on September 13, 2012.⁶ The Fed committed to buying \$40 billion of MBS and \$45 billion of longer-term US Treasuries per month for an indefinite period. After false expectations of a tapering in the amount of monthly purchases under QE3 in May 2013, the Fed announced seven consecutive reductions in the rate of asset purchases of \$10 billion per month between December 18, 2013 and September 17, 2014. When LSAPs were concluded in October 2014, the Fed held \$4.5 trillion of securities outright.⁷

The Fed have adopted numerous forms of forward guidance since December 2008 (Geraats, 2014). Initial guidance was qualitative, informing agents that the policy rate would be maintained “for some time” (December 2008) or “for an extended period” (March 2009). Subsequently, quantitative forward guidance was provided, including calendar-based guidance (August 2011) — informing agents that economic conditions were “likely to warrant exceptionally low levels for the federal funds rate” at least until a specified date — and threshold-based guidance (December 2012) — stating that the “exceptionally low range for the federal funds rate will be appropriate at least as long as the unemployment rate remains above 6.5 percent”.⁸ Moreover, the Fed engaged in forward guidance with respect to QE3, determining the size, pace and composition of future purchases in relation to future economic conditions.

Understanding the relative importance of different transmission channels of unconventional monetary policy is important because it can inform current and future policy. This study is motivated by the differing policy implications of the signalling and portfolio balance channels. Unconventional monetary policy can have *signalling* effects by influencing agents’ expectations of the future policy rate path. Forward guidance can do this directly, though LSAPs can have signalling benefits if they are perceived to signal a lower policy rate path for longer, especially when announced in advance of actual purchases. A policy that works through the signalling channel is likely to be most effective when it is clearly communicated, such that private sector expectations react to it. Portfolio rebalancing can occur as a result of LSAPs. By purchasing a longer-term asset from secondary markets, the central bank reduces the supply available to investors, bidding up the price and reducing the yield of the asset. With lower returns on their remaining holdings of the asset, investors can rebalance their portfolio to seek higher returns,

⁴www.federalreserve.gov/newsevents/speech/bernanke20100827a.htm.

⁵www.federalreserve.gov/newsevents/speech/bernanke20101015a.htm.

⁶Between QE2 and QE3, the Fed also initiated a maturity extension program (MEP), which was announced in September 2011 and was concluded in late 2012. The MEP was designed to extend the average maturity of the Treasuries in the Fed’s portfolio, placing downward pressure on longer-term interest rates to support economic conditions. The Fed sold a total of \$667 billion shorter-term Treasury securities under the MEP, buying longer-term Treasuries with the proceeds.

⁷Federal Reserve Statistical Release H.4.1, Factors Affecting Reserve Balances.

⁸These quotes are available here: www.federalreserve.gov/newsevents/press/all/XXXXall.htm, where XXXX denotes the year.

demanding other assets. This readjustment will increase the prices and reduce the yields of other assets. The efficacy of the portfolio balance channel relies on the ‘large-scale’ of LSAPs to generate sufficient portfolio adjustment to reduce long-term rates. Moreover, its benefits are likely to be greatest when markets are not functioning normally (Vayanos and Vila, 2009).

Two key predictions of the interest rate channel have been tested thoroughly: (i) that LSAPs and forward guidance reduced longer-term interest rates on announcement dates; and (ii) that they had expansionary effects on output and inflation. Many authors have shown that LSAPs did reduce longer-term interest rates (e.g. Krishnamurthy and Vissing-Jorgensen, 2011; Gagnon, Raskin, Remache, and Sack, 2011; Christensen and Rudebusch, 2012; Wright, 2012; Bauer and Rudebusch, 2014) on a range of assets, not only those purchased (D’Amico and King, 2012; Rogers, Scotti, and Wright, 2014).⁹ Studies have shown that LSAPs averted deflation and provided an expansionary impulse for output (e.g. Baumeister and Benati, 2013; Gambacorta, Hofmann, and Peersman, 2012),¹⁰ though only IMF (2013) and Lloyd (2013) have explicitly considered the importance of the signalling and portfolio balance channels for the real economy.

The majority of existing work has assessed the relative importance of the channels against their *financial market effects*. Using decompositions of long-term interest rates, authors have linked the portfolio balance channel to the term premium and the signalling channel to estimated risk-neutral yields. Disagreement in the results from this literature originates from the different yield curve decompositions used. Gagnon et al. (2011) use the survey-augmented affine Gaussian arbitrage-free dynamic term structure model (GADTSM) decomposition of US Treasury yields by Kim and Wright (2005) to show that, in terms of the yield effect, the portfolio balance channel has dominated the signalling benefits of US LSAPs at the 10-year horizon. However, Christensen and Rudebusch (2012) and Bauer and Rudebusch (2014), using alternative decompositions of US Treasury yields, have found the opposite result.

In this chapter, I focus on the relative importance of the two channels in terms of their effects on the *real economy*, the ultimate goal of US policy. Using a structural vector autoregression (SVAR) methodology, Lloyd (2013) concludes that the signalling channel was relatively more effective than the portfolio balance channel for US real GDP growth using the Kim and Wright (2005) decomposition of the 10-year Treasury yield. This chapter’s findings extend and reinforce those in Lloyd (2013). To reach this conclusion, I offer novel solutions to two challenges in existing literature: (i) the yield curve decomposition associated with signalling and portfolio rebalancing; and (ii) the identification of signalling and portfolio balance shocks to the macroeconomy.

In response to the first challenge, I compare three decompositions of the nominal US Treasury yield curve into risk-neutral yields and term premia:¹¹ (i) a bias-corrected model (Bauer, Rudebusch, and Wu, 2012), used to assess LSAPs by Bauer and Rudebusch (2014); (ii) a survey-augmented model (Kim and Wright, 2005), used to assess LSAPs by Gagnon et al. (2011);¹²

⁹Joyce, Lasaosa, Stevens, and Tong (2011) reach similar conclusions for the UK.

¹⁰Kapetanios, Mumtaz, Stevens, and Theodoridis (2012) provide similar results for the UK.

¹¹In GADTSMs, the risk-neutral yield corresponds to the average expected future short-term interest rate path, plus a convexity (Jensen’s inequality) term. This convexity term is, in practice, small, and so the terms ‘risk-neutral yield’ and ‘average expected future short-term interest rate’ are used interchangeably in the literature and this chapter.

¹²I use the survey-augmented model estimated by Kim and Wright (2005) estimated using the algorithm of Kim and Orphanides (2012), first circulated in Kim and Orphanides (2005), as opposed to the survey-augmented model estimated using the algorithm of Guimarães (2014), because the former of these is used by Gagnon et al. (2011), perhaps the most widely referenced US LSAP event study to date.

and (iii) the overnight indexed swap (OIS) rate-augmented decomposition proposed in chapter 3. As chapter 3 details, GADTSMs suffer from an identification problem that results in estimates of interest rate expectations that are spuriously stable. In this chapter, I provide complementary evidence to chapter 3 and show that, in comparison to financial market-based and survey expectations of future short-term interest rates, the interest rate expectation estimates from OIS-augmented models are superior to estimates from the bias-corrected and survey-augmented GADTSMs. Moreover, I document that, despite efforts to improve identification, the survey-augmented Kim and Wright (2005) decomposition does not fully overcome the identification problem and attributes too much variation in interest rates to term premia. Thus, the financial market event study estimates with the Kim and Wright (2005) decomposition represent a lower bound for the relative efficacy of the signalling channel.

Using these yield curve decompositions, I carry out an event study of the financial market effects of signalling and portfolio rebalancing. This is the first application of the OIS-augmented GADTSM to the analysis of macroeconomic policy at the ELB. I find that LSAP and forward guidance announcements had sizeable effects on interest rates and their associated expectations and term premia components. At the 2, 5 and 10-year horizons, the OIS-augmented model attributes 74.19-95.29% of the decline in yields on announcement days to signalling. Interestingly, interest rate expectations were affected well beyond the 2-year horizon traditionally associated with monetary policy's transmission lags. The results from the OIS-augmented decomposition differ starkly from the corresponding results using the survey-augmented Kim and Wright (2005) decomposition, which attributes only 22.85-30.41% of event-day yield declines to signalling at the 2, 5 and 10-year horizons.

To tackle the second challenge and estimate the relative effects of signalling and portfolio rebalancing on real economic outcomes, I set up an SVAR. The baseline VAR includes four monthly variables: industrial production, the consumer price index, and the risk-neutral yield and term premium components of longer-term interest rates derived from each of the three yield curve decompositions in turn. I identify structural shocks using combinations of zero-impact and sign restrictions, with signalling shocks propagating through the risk-neutral yield and portfolio rebalancing shocks through the term premium. To separately identify the signalling and portfolio balance shocks, I assess the robustness of results to two restriction schemes. In the first scheme, I impose that portfolio balance shocks cannot contemporaneously affect the risk-neutral component of long-term rates — this captures a 'pure term premium' shock. In the second, I impose the opposite: signalling shocks cannot immediately effect the term premium — this captures a 'pure expectations' shock.

Under both restriction schemes I find that an expansionary signalling shock has significantly positive lagged effects on US industrial production and consumer prices. In contrast, an expansionary portfolio rebalancing shock has insignificant effects on these two variables, indicating that signalling exerted a more powerful effect on US industrial production and consumer prices than portfolio rebalancing. The signalling shock explains around two-thirds to three-quarters of the total peak industrial production increase due to the long-term interest rate shocks. The results are robust to: (i) the inclusion of bank credit and the real exchange rate as controls; (ii) the interest rate maturity considered — even with longer-maturity yields, which place greater weight on portfolio rebalancing in the event study, the signalling channel is shown to be relatively more important for real outcomes; (iii) the sample length; and (iv) the term structure

decomposition used — the result even holds in SVAR specifications using the Kim and Wright (2005) yield curve decomposition, which places the lowest weight on signalling in the event study. The results are also robust to using the 2-year OIS rate, instead of the 2-year risk-neutral yield, alongside the 2-year term premium in the VAR to account for the possibility that, because both the risk-neutral yield and the term premium are estimated within the same GADSTM, they do not vary independently. My results suggest that current and future unconventional monetary policy action by the Fed may reap greater economic rewards if combined with clear communication about the future short-term interest rate path.

The remainder of this chapter is structured as follows. The transmission channels of unconventional monetary policy are defined in section 4.2. Section 4.3 presents the yield curve decompositions used in the event study (section 4.4) and SVARs (section 4.5). Section 4.6 concludes.

4.2 Transmission Channels

Unconventional monetary policy can affect the economy via numerous channels. The interest rate channel is the primary focus of this analysis. By purchasing assets directly from secondary markets, central bank LSAPs can raise asset prices and reduce a range of interest rates that investors face. This can positively impact upon the real economy through, *inter alia*, reduced borrowing costs and positive wealth effects.

A sizeable literature has amassed discussing a number of interest rate mechanisms through which LSAPs affect the real economy (Krishnamurthy and Vissing-Jorgensen, 2011). The focus of my study, in line with work by Gagnon et al. (2011), Bauer and Rudebusch (2014) and others, is on two components of the interest rate channel: signalling and portfolio rebalancing.

Primarily, I consider these channels because of their link to the canonical decomposition of the long-term interest rate into a risk-neutral expected future short-term interest rate and term premium component:

$$y_{L,t} = \frac{1}{L} \mathbb{E}_t \sum_{l=0}^{L-1} y_{1,t+l} + tp_{L,t} \quad (4.1)$$

where $y_{L,t}$ is the L -period government bond yield at time t , $y_{1,t}$ is the one-period (net) interest rate and $tp_{L,t}$ is the L -period term premium. In line with the existing literature, I link signalling to the risk-neutral component and portfolio rebalancing to the term premium.

Additionally, the signalling and portfolio balance channels are of direct relevance to policy. They make up the language of policymakers. Bernanke (2010) emphasised the importance of the portfolio balance channel as a means through which LSAPs can affect the economy:

“I see the evidence as most favorable to the view that such purchases work primarily through the so-called portfolio balance channel...”

Moreover, the policy implications relevant to each of the two channels differ: the signalling channel implies that policymakers should clearly communicate the future path of short-term interest rates; the portfolio balance channel relies on the stock of assets purchased being sufficient enough in scale to influence term premia.

4.2.1 Signalling

The signalling channel refers to any effect that (unconventional) monetary policy announcements have on investors' expectations of future short-term policy rates. Such expectations can be influenced by future macroeconomic outcomes or the expected conduct of monetary policy. If, following forward guidance or LSAP announcements, investors anticipate the central bank to keep interest rates lower for longer, then the announcement will influence long-term interest rates by reducing expected future short-term interest rates. Although this definition subsumes signals about future policy rates from forward guidance or LSAPs, it excludes any LSAP policy anticipation or announcement effects that cause immediate portfolio changes. These will be attributed to the term premium. In addition, although the main purpose of forward guidance is to influence expectations of future policy rates, this could also reduce uncertainty about future interest rates and thereby term premia (Akkaya, 2014).

In 2008, when the Fed initiated LSAPs, many critiques of the policy cited irrelevance and neutrality propositions (e.g. Wallace, 1981). Integral to all of these results are assumptions regarding: timing; household homogeneity; perfect asset substitutability; non-distortionary taxation; and the link between government and central bank balance sheets. The logic behind these results is as follows. The purchase of long-term assets by the central bank can increase households' *pre-tax* state-contingent income. However, the purchase of the asset does not remove risk from the aggregate economy. LSAPs will reduce the returns earned by the central bank portfolio, necessitating an increase in lump-sum taxation by a non-distortionary government to balance the joint government and central bank budget constraint. The *after-tax* state-contingent income of homogeneous households will be unchanged, rendering LSAPs neutral for the economy. Within these models, LSAPs can only circumvent neutrality propositions through signalling; portfolio rebalancing is ineffective. In Eggertsson and Woodford (2003), only when LSAPs are *perceived* to engender a commitment to keep interest rates lower for longer can they stimulate the real economy. Bhattarai, Eggertsson, and Gafarov (2015) show that, following the purchase of longer-term assets and a shortening of the duration of privately held outstanding government debt, it is optimal for the central bank to keep short-term interest rates lower for longer to avoid capital losses on their balance sheet. Therefore, at the ELB, LSAPs can optimally stimulate the real economy by lowering the expected future path of real short-term interest rates.

4.2.2 Portfolio Rebalancing

The portfolio balance channel is linked to movements in term premia. By purchasing longer-term assets from the private sector, LSAPs concurrently increase the private sectors' holdings of short-term reserves. For investors who view different asset classes and maturities as imperfect substitutes to willingly accept this change, the price of longer-term assets must rise and their yield fall. To the extent that this change occurs independently of the short-term interest rate, it works through the term premium on longer-term assets. With lower long-term asset returns, investors will rebalance their portfolios, searching for higher yields by demanding other longer-term assets. This demand-driven rebalancing will inflate prices and reduce term premia on a range of long-term assets. D'Amico and King (2012) show that, although the term premia reduction was largest for the assets purchased by the Fed, US LSAPs did engineer declines in the term premia on a range of other longer-term assets. Ultimately, the lower term premia and

higher asset prices that result from portfolio rebalancing can transmit to the real economy by reducing borrowing costs for the private sector and generating positive wealth effects for private asset holders. The strength of portfolio rebalancing depends on the stock of assets purchased.

Because the term premium is defined to include compensation for interest rate risk, forward guidance may also affect term premia. If, following central bank announcements, investors' uncertainty surrounding the future path of short-term interest rates falls, this will be reflected in lower term premia.¹³ Similarly, forward guidance about LSAPs can be expected to influence term premia by instigating portfolio changes on announcement days.

Irrelevance propositions preclude portfolio rebalancing's efficacy in many macroeconomic models. To admit such effects, theorists have incorporated imperfect asset substitutability (Tobin, 1956, 1969) with agent heterogeneity. Harrison (2011, 2012) and Chen, Cúrdia, and Ferrero (2012) show that LSAPs can benefit the real economy via portfolio rebalancing within theoretical models.

4.2.3 Other Channels

Unconventional policy can transmit to the real economy through other channels. Joyce, Miles, Scott, and Vayanos (2012) discuss a credit channel through which LSAPs can affect output and inflation, independent of long-term interest rates. By purchasing assets from non-bank financial institutions, the deposits these institutions place in banks may rise. If deposits exceed banks' demand for liquidity, banks may be more willing to extend credit in the form of lending or less willing to contract it if they suffer funding losses from other sources. This channel is likely to be most effective when bank funding is dysfunctional, as it was after the 2007-2008 financial crisis. It is likely to be relevant for US LSAPs, where the Fed has purchased the majority of its assets from households (including hedge funds), broker dealers and insurance companies (Carpenter, Demiralp, Ihrig, and Klee, 2013).

Unconventional monetary policy may also have international effects, through an exchange rate channel. If forward guidance or LSAP announcements reduce contemporaneous interest rates and expected future rates, they may lead international investors to seek higher returns away from the domestic economy. Theoretically, this should depreciate the domestic currency, *ceteris paribus*, aiding the price competitiveness of exports and, thus, domestic output. Bauer and Neely (2012) argue that since these changes work through long-term interest rates, these international effects are due to signalling and portfolio rebalancing. As a result, I assess the relative importance of signalling and portfolio rebalancing both with and without controls for the international transmission of policy in section 4.5.

4.3 Decompositions of the Yield Curve

To assess the relative importance of the signalling and portfolio rebalancing channels, I rely on decompositions of the yield curve, informed by (4.1), into risk-neutral yields (expectations of future short-term interest rates) and term premia. Like other authors (e.g. Gagnon et al., 2011; Bauer and Rudebusch, 2014), I associate the signalling channel with the risk-neutral yields

¹³The term premium may also include liquidity premia. In my study, any liquidity effects due to unconventional monetary policy are attributed to portfolio rebalancing.

and portfolio rebalancing with the term premium. To decompose yields, I estimate three no-arbitrage Gaussian affine dynamic term structure models (GADTSMs) and compare the results across different models. The differing conclusions in the existing literature are driven by the different GADTSMs used. For instance, Gagnon et al. (2011) use the survey-augmented Kim and Wright (2005) GADTSM and conclude that the effects of portfolio rebalancing dominate those of signalling, while Bauer and Rudebusch (2014) use the bias-corrected Bauer et al. (2012) GADTSM and attribute a larger proportion of influence to signalling.

The differing predictions of GADTSMs in analyses of signalling and portfolio rebalancing arise from the identification problem discussed in chapter 3 that results in estimates of interest rate expectations that are spuriously stable (e.g. Kim and Orphanides, 2012; Guimarães, 2014). In response to this, three solutions have been proposed: bias correction (Bauer et al., 2012); survey-augmentation (Kim and Orphanides, 2012);¹⁴ and OIS-augmentation (chapter 3). In the following sub-sections, I describe the data and algorithms I use to estimate these three GADTSMs. I compare model-implied risk-neutral yields to comparable-horizon federal funds futures rates and survey expectations, and show that the OIS-augmented model provides superior estimates of interest rate expectations for the period of relevance to this chapter.¹⁵

4.3.1 Estimation of GADTSMs

To foster the closest possible comparison to the related literature, I compare the OIS-augmented model to the survey-augmented (Kim and Wright, 2005) and bias-corrected (Bauer et al., 2012) GADTSMs. The Kim and Wright (2005) data I use is publicly available and estimated with daily frequency bond yield data from July 18, 1990 to December 31, 2015 with 3 and 6-month T-Bill yields and 1, 2, 4, 7 and 10-year US Treasury zero-coupon bond yields (Gürkaynak, Sack, and Wright, 2007a).¹⁶ I estimate the bias-corrected model with the same data, using the algorithm of Bauer et al. (2012, Section 4). Because US OIS rate data is only available from late-2001, I estimate the OIS-augmented decomposition using data from January 2, 2002 to December 31, 2015 to isolate the benefits of OIS-augmentation, and using the same bond maturities as in chapter 3.¹⁷ I use 3, 6, 12 and 24-month OIS rates, as in the 4-OIS-augmented model in chapter 3 (hereafter the ‘OIS-augmented decomposition’). Chapter 3 documents that this model risk-neutral yields that are superior to those from existing GADTSMs, excluding Kim and Wright (2005), for the 2002-2016 period, exhibiting the lowest root mean square error (RMSE) fit *vis-à-vis* federal funds futures rates and survey expectations. For each model, three pricing factors determine bond prices.¹⁸

Figure 4.1 presents the results from the three GADTSMs at the 2-year horizon. Panel A plots the actual time series of the 2-year yield against the fitted values from the three GADTSMs over the 2002-2015 period. The illustration corroborates an important finding in chapter 3: GADTSM-augmentation does not compromise the overall fit of the model with respect to actual

¹⁴The Kim and Wright (2005) decomposition is estimated using the Kim and Orphanides (2012) algorithm, first circulated in Kim and Orphanides (2005).

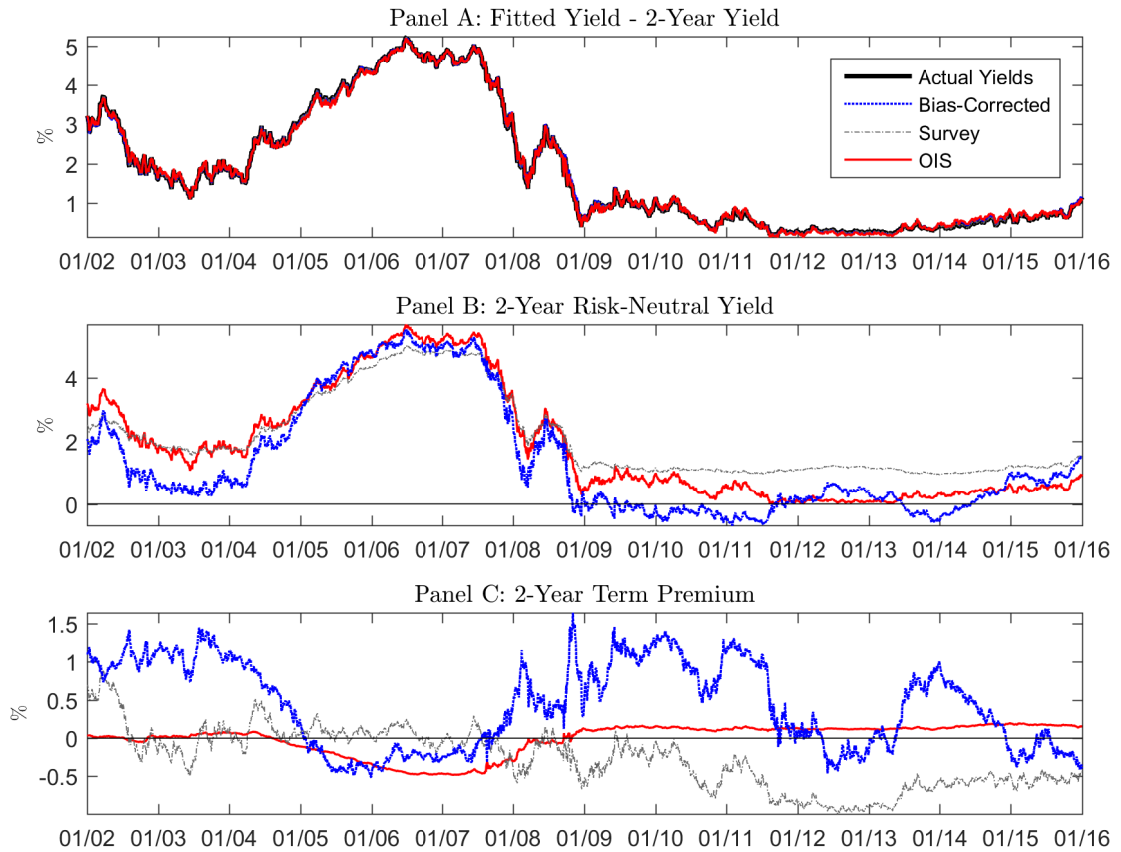
¹⁵All of these models are linear, so the results should be interpreted with some caution in light of the ELB that constrained short-term policy rates from 2008 onwards. Nevertheless, the differing ability of the three models in matching measures of interest rate expectations is striking.

¹⁶T-Bill rates are converted from the discount to the yield basis. Data sources are listed in appendix A.1.

¹⁷3 and 6-month T-Bill rates and 12, 18, ..., 60, 84, and 120-month zero-coupon Treasury bond yields.

¹⁸Kim and Wright (2005) and Bauer et al. (2012) also use three pricing factors. Litterman and Scheinkman (1991) demonstrate that the first three principal components of bond yields explain over 95% of their variation.

Figure 4.1: Estimated Yield Curve Decomposition: July 1990-December 2015



Note: In panel A, I plot the actual 2-year bond yield and fitted 2-year bond yields from each of three GADTSMs. In panels B and C, I plot the estimated risk-neutral yields and term premia from the three GADTSMs, respectively. The three models are: (i) the bias-corrected model of Bauer et al. (2012) (Bias-Corrected); (ii) the survey-augmented model of Kim and Wright (2005) (Survey); and (iii) the OIS-augmented model of chapter 3 (OIS). The bias-corrected and survey-augmented models are estimated using daily data from July 18, 1990 to December 31, 2015. The OIS-augmented model is estimated using daily data from January 2, 2002 to December 31, 2015. All models use three pricing factors. All figures are in annualised percentage points.

bond yields. The series co-move extremely closely.¹⁹ As stated in chapter 3, this finding is intuitive. Survey and OIS-augmentation have been proposed to improve the identification of risk-neutral yields. Even in unaugmented GADTSMs, bond yield data is sufficient for the accurate fitting of actual bond yields.

4.3.2 Interest Rate Expectations

Unlike fitted yields, panels B and C of figure 4.1 illustrate that the risk-neutral yields and term premia from each of the GADTSMs differ markedly. The differences are a direct consequence of the identification problem. The risk-neutral yields differ starkly from late-2008 onwards, the period most relevant to this analysis. The 2-year risk-neutral yield from the survey-augmented Kim and Wright (2005) decomposition remains persistently above 1% from December 2008 onwards, while the bias-corrected and OIS-augmented models attribute a greater proportion of the fall in yields during 2008 to falling expectations of future short-term interest rates. In fact,

¹⁹The residuals of the fitted yields are extremely similar across models at all maturities.

the 2-year risk-neutral yield from the bias-corrected model is persistently negative from mid-2009 to late-2011, counter-factually implying that investors' average expectation of future short-term interest rates was negative. In contrast, the 2-year risk-neutral yield from the survey-augmented and OIS-augmented models never fall negative.²⁰

In figure 4.1, as in chapter 3, the 2-year term premium from the OIS-augmented model is persistently negative from mid-2004 to mid-2008. This is a direct consequence of the accurate fitting of risk-neutral yields. However, this feature is not true for all maturities; estimated term premia at longer horizons are frequently and persistently positive.

To accurately attribute yield changes to signalling and portfolio rebalancing effects, it is necessary to attain accurate measures of the risk-neutral yields and term premia used to identify the channels. With this goal in mind, I compare the model-implied interest rate expectations, from the risk-neutral yields, to federal funds futures rates and survey expectations. The preferred GADTSM should accurately reflect the qualitative and quantitative evolution of comparable-horizon survey and market-implied expectations.

Risk-Neutral Yields and Federal Funds Futures Rates

I first compare the GADTSM-implied risk-neutral yields to federal funds futures (FFFs) rates, by performing the following steps.²¹ First, I construct a FFFs-implied expectation of the average short-term interest rate over a 12-month period. To do this, I calculate the arithmetic mean of the 1, 2, ..., 12-month ahead FFFs rates on the final day of each calendar month, generating a monthly frequency series of average market-implied interest rate expectations over the subsequent 12-months.²² Second, I compare the monthly frequency FFF-implied expectation to the corresponding-horizon 1-year risk-neutral yield from each GADTSM on the final day of each calendar month.

Table 4.1 reports a root mean square error (RMSE) comparison of the FFF-implied and GADTSM-implied expectations for three sample periods: January 2002 to December 2015; the baseline SVAR sample period from November 2008 to April 2013; and November 2008 to December 2015. On a RMSE basis, the OIS-augmented model unambiguously provides superior estimates of expected future short-term interest rates, as measured by FFFs rates. Between November 2008 and December 2015, the RMSE of the OIS-augmented model approximately half of the RMSE of the survey-augmented model, and almost a third of the RMSE of the bias-corrected model. For all three periods, the bias-corrected model provides the worst fit of FFF-implied market expectations.

²⁰As in chapter 3 this is true at all horizons for the OIS-augmented model, despite the fact that additional restrictions are not imposed on the model to prevent interest rate expectations from going negative. This represents an important contribution in light of recent computationally burdensome proposals for term structure modelling at the ELB (see, for example, Christensen and Rudebusch, 2013a,b).

²¹Ideally, I would follow the steps in section 3.6.2, and compare FFFs rates to model-implied risk-neutral forward yields 1, 2, ..., 11 months ahead. However, because I do not estimate the survey-augmented Kim and Wright (2005) decomposition, and because they do not report 1, 2, ..., 11-month risk-neutral yields, I must alter the analysis. In chapter 3, I show that the OIS-augmented model performs unambiguously better than other models at the 1, 2, ..., 11 month horizons.

²²I only use FFFs rates on the final day of each calendar month due to the maturity structure of FFFs contracts (see chapter 2). An n -month contract traded on day t_i of the calendar month t has the same settlement period as an n -month contract traded on a different day t_j in the same calendar month t . For this reason, the horizon of the FFFs-implied expectation and the 1-year risk-neutral yield only align on the final day of each calendar month. I use the arithmetic mean in accordance with FFFs market convention — see CME Rulebook, Chapter 22, 22101: www.cmegroup.com/rulebook/CBOT/V/22/22.pdf.

Table 4.1: GADTSM-Implied Expectations: Root Mean Square Error (RMSE) of the 1-Year Risk-Neutral Yield *vis-à-vis* the Federal Funds Futures-Implied 1-Year Expectation

Model	RMSE vs. 1-Year FFF-Implied Expectation		
	Sample		
	Jan. 2002 to Dec. 2015	Nov. 2008 to Apr. 2013	Nov. 2008 to Dec. 2015
Bias-Corrected	0.5275	0.5386	0.4707
Survey-Augmented	0.3462	0.3521	0.3456
OIS-Augmented	0.2522	0.2138	0.1713

Note: RMSE of the 1-year risk-neutral yields from each of the three GADTSMs in comparison to the federal funds futures-implied expectation. The three models are: (i) the bias-corrected model (Bauer et al., 2012); (ii) the survey-augmented model (Kim and Wright, 2005); and (iii) the OIS-augmented model (Lloyd, 2017a). The bias-corrected and survey-augmented models are estimated using daily data from July 18, 1990 to December 31, 2015. The OIS-augmented model is estimated using daily data from January 2, 2002 to December 31, 2015. All models use three pricing factors. All figures are in annualised percentage points. The lowest RMSE model has been emboldened for ease of reading.

Risk-Neutral Yields and Survey Expectations

I also compare the GADTSM-implied interest rate expectations to survey expectations. Because survey expectations reflect respondents' expectations of future short-term interest rates, the risk-neutral yields from the preferred GADTSM should closely align with corresponding-maturity survey expectations, especially during the post-2008 period of interest.

Formally, I compare the estimated 6-month and 1-year risk-neutral yields to corresponding-horizon average short-term interest rate expectations from surveys.²³ I calculate approximate future short-term interest rate expectations using data from the quarterly *Survey of Professional Forecasters* at the Federal Reserve Bank of Philadelphia, using the weighted geometric approximations detailed in appendix A.2. To compare the estimated risk-neutral yields to these survey expectations, I calculate the RMSE of the risk-neutral yield *vis-à-vis* the corresponding horizon survey expectation on survey submission deadline dates.

Table 4.2 presents the numerical results for the comparison of 6 and 12-month expectations. The results indicate that the OIS-augmented GADTSM unambiguously provides the best fit for survey expectations. Of particular note is the performance of the OIS-augmented model over the baseline 2008 Q4 to 2013 Q2 sample most relevant to the subsequent analysis. Here the RMSE fit of the 1-year survey expectation from the OIS-augmented model is almost one-third of the RMSE fit of the survey-augmented model and around a quarter of the RMSE fit of the bias-corrected model.

Overall, the OIS-augmented decomposition provides superior estimates of interest rate expectations, in comparison to both FFFs rates and survey expectations. Hereafter, the OIS-augmented model is deemed the preferred model of interest rate expectations.

²³Estimates of the Kim and Wright (2005) decomposition are only available for 1-year bond maturities, or more, so I am unable to compare this to survey forecasts at the 6-month horizon. Nevertheless, I compare the bias-corrected and OIS-augmented models at the 6-month horizon, to stress the superiority of OIS-augmentation over the bias-corrected model at multiple horizons.

Table 4.2: GADTSM-Implied Expectations: Root Mean Square Error (RMSE) of the In-Sample Risk-Neutral Yields *vis-à-vis* 6-Month and 1-Year Survey Expectations

Model	Sample		
	2002 Q1 to 2015 Q4	2008 Q4 to 2013 Q2	2008 Q4 to 2015 Q4
	RMSE vs. 6-Month Survey Expectation		
Bias-Corrected	0.2725	0.2977	0.2606
Survey-Augmented	N/A	N/A	N/A
OIS-Augmented	0.1505	0.1132	0.0979
	RMSE vs. 1-Year Survey Expectation		
Bias-Corrected	0.4555	0.5016	0.4329
Survey-Augmented	0.2992	0.3519	0.3292
OIS-Augmented	0.1685	0.1400	0.1306

Note: RMSE of the risk-neutral yields from each of the three GADTSMs in comparison to approximated survey expectations. The three models are: (i) the bias-corrected model (Bauer et al., 2012); (ii) the survey-augmented model (Kim and Wright, 2005); and (iii) the OIS-augmented model (Lloyd, 2017a). The bias-corrected and survey-augmented models are estimated using daily data from July 18, 1990 to December 31, 2015. The OIS-augmented model is estimated using daily data from January 2, 2002 to December 31, 2015. All models use three pricing factors. All figures are in annualised percentage points. The lowest RMSE model at each maturity, for each sub-sample, has been emboldened for each of reading.

4.4 Financial Market Impact of LSAPs and Forward Guidance

Before assessing the effect of shocks to longer-term interest rates on real economic outcomes, I document the impact of LSAP and forward guidance announcements on interest rates by carrying out an event study. To label the shocks identified in section 4.5 as the ‘signalling’ and ‘portfolio rebalancing’ effects of LSAPs and forward guidance, policy announcements must have exerted a significant impact on the risk-neutral yield and term premium components of longer-term bond yields. I verify this here.

Event studies are ubiquitous in the literature assessing the financial market effects of unconventional monetary policy. As is the norm, I evaluate the change in interest rates within a one-day event window on event days where notable announcements pertaining to forward guidance or the *expansion* of LSAPs occurred.

Event studies rely on the lumpy nature of monetary policy announcements. Although US unconventional monetary policy announcements have occurred at different points in time and at irregular intervals, they have been multifaceted and have become increasingly complex.²⁴ Numerous policies have been announced in a single statement. For instance, on December 16, 2008, the Fed announced that the target range for the federal funds rate would be reduced to 0-0.25%, that interest rates would be kept low “for some time”, and that LSAPs would be continued. Thus, I study forward guidance and LSAPs jointly, as on some event days the effects of the two are not separately identifiable.

My results extend upon the existing US unconventional monetary policy event study literature in three ways. First, this is the first study to apply the OIS-augmented yield curve decomposition of chapter 3 to the analysis of macroeconomic policy at the ELB. Second, I consider a longer sample period of events: from November 25, 2008 to April 2013, though the

²⁴Rogers, Scotti, and Wright (2014) note that the average word count for FOMC statements has increased from around 200 words in 2008 to over 600 in 2013.

last event date corresponding to expansionary monetary policy is December 12, 2012. Third, the classification of events differs to those of Rogers, Scotti, and Wright (2014) and Gilchrist, López-Salido, and Zakrajsek (2015), who are the only other authors to explicitly consider the simultaneous occurrence of forward guidance and LSAP announcements on event days. These authors classify the events in a binary manner, as either predominantly LSAP-related or forward guidance-related. I add a third classification to account for event dates on which both notable LSAP and forward guidance events took place. Admitting the joint impact of these policies is especially important, as either policy is likely to contaminate event studies into the other. For instance, in an LSAP-only event study, the sizeable reduction in US Treasury yields on March 18, 2009 may be entirely attributed to the announced purchase of longer-term Treasuries as part of QE1. However, on the same day the Fed altered its forward guidance from stating that it would maintain the policy rate at its lower bound “for some time” to “an extended period”. By defining this as a combined LSAP and forward guidance event, I explicitly capture the multifaceted nature of unconventional monetary policy. Finally, I consider movements in the risk-neutral yield and term premium at multiple horizons: specifically 2, 5 and 10 years. Although movements in yields of different maturities are highly correlated, there is no *a priori* reason to expect changes in interest rate expectations to be equally important at all time horizons. In fact, as forward guidance is often strongly linked — either explicitly when time-dependent, or implicitly otherwise — to a 1 to 2-year horizon, it seems likely that signalling effects will be most important at these tenors. Additionally, 10-year interest rate movements, which are the sole focus of some existing LSAP event studies, may not be the most relevant for economic activity. By additionally considering 2 and 5-year rates, I am better able to account for heterogeneous effects of signalling and portfolio rebalancing across horizons.

Table 4.3 presents the list of 16 announcement dates that I consider. All announcements are based on a set of official communications by the Fed and speeches by senior Fed officials, which contained new information on unconventional policy. To select the events, I independently scoured all Fed press releases.²⁵ To be included in the event set, the news had to mark a notable, broadly unanticipated change in LSAP or forward guidance policy. Many of the events in the first half of the study corroborate with those in other event studies (Gagnon et al., 2011; Christensen and Rudebusch, 2012; Filardo and Hofmann, 2014).

4.4.1 Model-Free Evidence

I first consider the behaviour of Treasury yields and OIS rates on event days without a model for investors’ interest rate expectations. This ‘model-free’ evidence is a useful benchmark for comparing the results attained using formal yield curve decompositions.

OIS rates are particularly relevant for studies of the signalling channel of unconventional monetary policy, as they are associated with investors’ expectations of the future short-term interest rate path. Chapter 2 explains how OIS rates theoretically reflect interest rate expectations and demonstrates that, for the 2002-2016 period, US OIS contracts out to the 2-year horizon provide accurate information about investors’ expectations of future short-term interest rates. Longer-maturity OIS rates include term premia that are increasing in contract maturity, which mean this model-free evidence can only provide illustrative evidence on the importance

²⁵These press releases are available here: www.federalreserve.gov/newsevents/press/all/XXXXall.htm, where XXXX should be replaced by the year of interest.

Table 4.3: Event Set and Descriptions

#	Date	Description
I	25/11/2008	Initial LSAP announcement: <i>LSAPs</i> - Fed to purchase MBS and agency bonds. ^{GRRS,CR,W}
II	01/12/2008	Bernanke speech: <i>LSAPs</i> - US Treasuries <i>may</i> be purchased. ^{GRRS,CR,W,a}
III	16/12/2008	FOMC statement: <i>FG</i> - "... weak economic conditions are likely to warrant exceptionally low levels of the federal funds rate for some time"; <i>LSAPs</i> - Further mention of possible Treasury purchases. ^{GRRS,CR,W,FH}
IV	28/01/2009	FOMC statement: <i>LSAPs</i> - "The Federal Reserve continues to purchase large quantities of agency debt and mortgage-backed securities to provide support to the mortgage and housing markets, and it stands ready to expand the quantity of such purchases and the duration of the purchase program as conditions warrant". ^{GRRS,CR,W}
V	18/03/2009	FOMC statement: <i>FG</i> - "... economic conditions are likely to warrant exceptionally low levels of the federal funds rate for an extended period"; <i>LSAPs</i> - Purchase of long-term Treasuries announced. ^{GRRS,CR,W,FH}
VI	10/08/2010	FOMC statement: <i>LSAPs</i> - Fed to reinvest holdings of assets purchased under 'QE1' to keep overall value of asset stock constant. ^W
VII	27/08/2010	Bernanke speech at Jackson Hole: <i>LSAPs</i> - "additional purchases ... would be effective". ^{W,b}
VIII	21/09/2010	FOMC statement: <i>LSAPs</i> - Fed to reinvest holdings of assets purchased. ^W
IX	15/10/2010	Bernanke speech at Boston Fed: <i>FG & LSAPs</i> - "the FOMC is prepared to provide additional accommodation if needed to support the economic recovery and to return inflation over time to levels consistent with our mandate". ^{W,c}
X	03/11/2010	FOMC statement: <i>LSAPs</i> - 'QE2' announced; \$600bn purchase of long-term Treasuries over six months.
XI	09/08/2011	FOMC statement: <i>FG</i> - "... economic conditions ... are likely to warrant exceptionally low levels for the federal funds rate at least through mid-2013"; <i>LSAPs</i> - Fed to reinvest holdings of assets purchased under 'QE1' and 'QE2' to keep overall value of asset stock constant. ^{W,FH}
XII	26/08/2011	Bernanke speech at Jackson Hole: <i>FG & LSAPs</i> - "[T]he [Fed] has a range of tools. ... We will continue to consider those". ^{W,d}
XIII	25/01/2012	FOMC statement: <i>FG</i> - "... economic conditions ... are likely to warrant exceptionally low levels for the federal funds rate at least through late 2014". ^{FH}
XIV	31/08/2012	Bernanke speech at Jackson Hole: <i>FG & LSAPs</i> - "nontraditional policy tools ... can continue to be effective". ^e
XV	13/09/2012	FOMC statement: <i>FG</i> - "... low levels for the federal funds rate are likely to be warranted at least through mid-2015"; <i>LSAPs</i> - 'QE3' announced. ^{FH}
XVI	12/12/2012	FOMC statement: <i>FG</i> - "... exceptionally low range for the federal funds rate will be appropriate at least as long as the unemployment rate remains above 6.5 percent, inflation between one and two years ahead is projected to be no more than half a percentage point above the Committee's 2 percent longer-run goal, and longer-term inflation expectations continue to be well anchored". ^{FH}

Abbreviations: LSAPs = Event date with LSAP news; FG = Event date with forward guidance announcement. ^{GRRS}: In the baseline event set of Gagnon et al. (2011); ^{CR}: Included in the event set of Christensen and Rudebusch (2012); ^W: Included in the important event set of Wright (2012); ^{FH}: Included in event set of Filardo and Hofmann (2014).

For speeches, the source is www.federalreserve.gov/newsevents/speech/XX where XX is [bernanke20081201a.htm](http://www.federalreserve.gov/newsevents/speech/bernanke20081201a.htm) for ^a; [bernanke20100827a.htm](http://www.federalreserve.gov/newsevents/speech/bernanke20100827a.htm) for ^b; [bernanke20101015a.htm](http://www.federalreserve.gov/newsevents/speech/bernanke20101015a.htm) for ^c; [bernanke20110826a.htm](http://www.federalreserve.gov/newsevents/speech/bernanke20110826a.htm) for ^d; and [bernanke20120831a.htm](http://www.federalreserve.gov/newsevents/speech/bernanke20120831a.htm) for ^e.

News source sampled from all Fed press releases 2008-2013: www.federalreserve.gov/newsevents/pressreleases.htm.

of signalling for the propagation of unconventional monetary policy.

I compare the changes in actual Treasury yields to comparable-horizon OIS rates on event days. For the 2-year OIS rate, quantitative similarities in the daily changes of OIS and Treasury rates indicate the existence of a common factor explaining the co-movement. Expectations of future short-term interest rates, which are reflected in both OIS and Treasury rates, are a prime candidate for this factor. However, because term premia are likely to contaminate daily changes in 5 and 10-year OIS rates, a close co-movement between OIS and Treasury rates at these horizons is less likely to be explained by a single common expectations factor.

Table 4.4 presents the results of the model-free event study, using US Treasury zero-coupon yields from Gürkaynak et al. (2007a) and OIS rates from *Bloomberg*. US unconventional monetary policy announcements did have a significant impact on financial markets at a range of horizons.²⁶ Treasury yields and OIS rates fell significantly on many of the event days. On the whole, the results depict a hump-shaped response of the Treasury yield curve in reaction to news: the largest cumulative fall in Treasury yields over the 16 announcement dates was at the 5-year maturity. The cumulative responses of OIS rates do not exhibit such a defined hump-shaped pattern across maturities. On LSAP-only and forward guidance-only event days, the cumulative fall in OIS rates is increasing in contract maturity. Though, on all event days, the 5-year OIS rate exhibits the largest cumulative fall, exceeding the 10-year figure by 0.75 basis points. The differing response of the OIS maturity structure to news, *vis-à-vis* Treasury yields, provides illustrative evidence of the term premia that exist in longer maturity OIS contracts.

The primary finding from table 4.4 is that the spread between 2-year Treasury and OIS rates moved very little on most event days. Changes in interest rate expectations are likely to have acted as a common factor driving this co-movement, indicating an important role for signalling effects in the transmission of US unconventional monetary policy to financial markets. Most strikingly, the total change in 2-year OIS rates on LSAP event days is 75.81% (−58.60 basis points) of the cumulative change in the 2-year Treasury yield (−77.30 basis points) on the same event days. The corresponding percentages for the forward guidance-only events and all events are 66.62% and 73.72% at the 2-year horizon. At the 5 and 10-year horizons the ratios of the cumulative falls in OIS rates to Treasury yields are higher than for the 2-year tenor. However, these figures are likely to reflect significant term premia within OIS contracts in addition to a common expectations factor, motivating the subsequent GADTSM-based study.

Treasury yields and OIS rates fell significantly on most event days.²⁷ However, on some event days, Treasury and OIS rates increased (e.g. IV, VII, IX, and XVI),²⁸ and their moves were statistically insignificant on other days (e.g. XII). Nevertheless, the conclusions are robust to the removal of these dates from the event set. When event days IV, VII, IX, XII and XVI are not included, the cumulative fall in the 2-year Treasury yield is 88.12 basis points, while the 2-year OIS rate cumulatively falls by 69.45 basis points. The 5 and 10-year Treasury yields fell by 166.57 and 164.30 basis points, respectively, on the remaining 11 event days, while the 5 and

²⁶To assess the significance of these daily changes, I control for other macroeconomic data releases. Formally, the tests assess whether the change on an event date is significantly different to the average change on a non-event day, controlling for macroeconomic data releases. See appendix C.1.2 for more details. Similar results are attained when the Citigroup Economic Surprise Index, a summary measure of differences between economic data releases (excluding monetary policy) and pre-announcement expectations, is used as a control variable.

²⁷Although yields fell by more on event XII than on other days (e.g. X), the statistical insignificance of the changes occurs because a preliminary US GDP release occurred on the same date. Preliminary US GDP releases also occurred on days I and VII, but the yield moves on these days were larger and, thus, statistically significant.

²⁸Appendix C.1.1 explains why Treasury yields increased on these event days with reference to news reports.

Table 4.4: One-Day Change in Actual US Treasury Yields and OIS Rates on Event Dates

#	Event Date & Description		Maturity - All Figures in Basis Points					
			2-Year		5-Year		10-Year	
			(1) Δyld_2	(2) Δi_2^{ois}	(3) Δyld_5	(4) Δi_5^{ois}	(5) Δyld_{10}	(6) Δi_{10}^{ois}
I	25/11/2008	L	-14.35***	-13.80***	-22.50***	-24.80***	-21.38***	-28.10***
II	01/12/2008	L	-11.86***	-12.80***	-21.42***	-20.70***	-21.55***	-19.10***
III	16/12/2008	F,L	-10.72***	-15.25***	-16.26***	-28.55***	-17.48***	-31.75***
IV	28/01/2009	L	+4.57***	+5.60***	+10.11***	+10.70***	+12.04***	+13.65***
V	18/03/2009	F,L	-26.41***	-12.00***	-47.08***	-26.70***	-51.88***	-37.70***
VI	10/08/2010	L	-2.69**	-1.20	-7.09***	-5.30***	-6.87***	-4.40***
VII	27/08/2010	L	+5.40***	+5.10***	+12.30***	+13.00***	+16.64***	+18.30***
VIII	21/09/2010	L	-3.71***	-4.05***	-9.57***	-10.30***	-10.73***	-12.40***
IX	15/10/2010	F,L	-1.21**	-0.50***	+2.61***	+1.80***	+8.62***	+4.30***
X	03/11/2010	L	-1.52***	-0.70***	-4.04***	-3.10***	+4.07***	-1.70***
XI	09/08/2011	F,L	-8.56***	-5.95***	-19.09***	-10.35***	-20.50***	-8.60***
XII	26/08/2011	F,L	-1.70	-0.75	-4.21	-4.65***	-3.50	-6.00
XIII	25/01/2012	F	-3.77***	-1.40***	-9.39***	-10.45***	-8.03***	-5.70***
XIV	31/08/2012	F,L	-3.67***	-1.10***	-6.43***	-7.00***	-7.02***	-7.10***
XV	13/09/2012	F,L	-0.86	-1.20	-3.70***	-4.80***	-2.93***	-5.60***
XVI	12/12/2012	F	+0.03	+0.25**	+2.25***	+2.65***	+5.71***	+4.10***
Total Change on Event Days								
<i>LSAP Events</i>			-77.30	-58.60	-136.37	-120.75	-122.47	-126.20
$(\Delta i^{ois} \div \Delta yld)$				(75.81%)		(88.55%)		(103.05%)
<i>FG Events</i>			-56.89	-37.90	-101.30	-88.05	-97.02	-94.05
$(\Delta i^{ois} \div \Delta yld)$				(66.62%)		(86.92%)		(96.94%)
<i>All Events</i>			-81.05	-59.75	-143.51	-128.55	-124.79	-127.80
$(\Delta i^{ois} \div \Delta yld)$				(73.72%)		(89.58%)		(102.41%)

Notes: L = Event date with LSAP news; F = Event date with forward guidance announcement; Δyld_n = Change in actual n -year zero-coupon bond yield on event days; Δi_n^{ois} : Change in actual n -year OIS rate on event days. Tests to determine the significance of daily changes are described in appendix C.1.2; t -statistics are calculated using Newey and West (1987) standard errors. Daily changes that are significant at the 1%, 5% and 10% levels are denoted with asterisks ***, ** and * respectively. *Data Sources*: Appendix A.1.

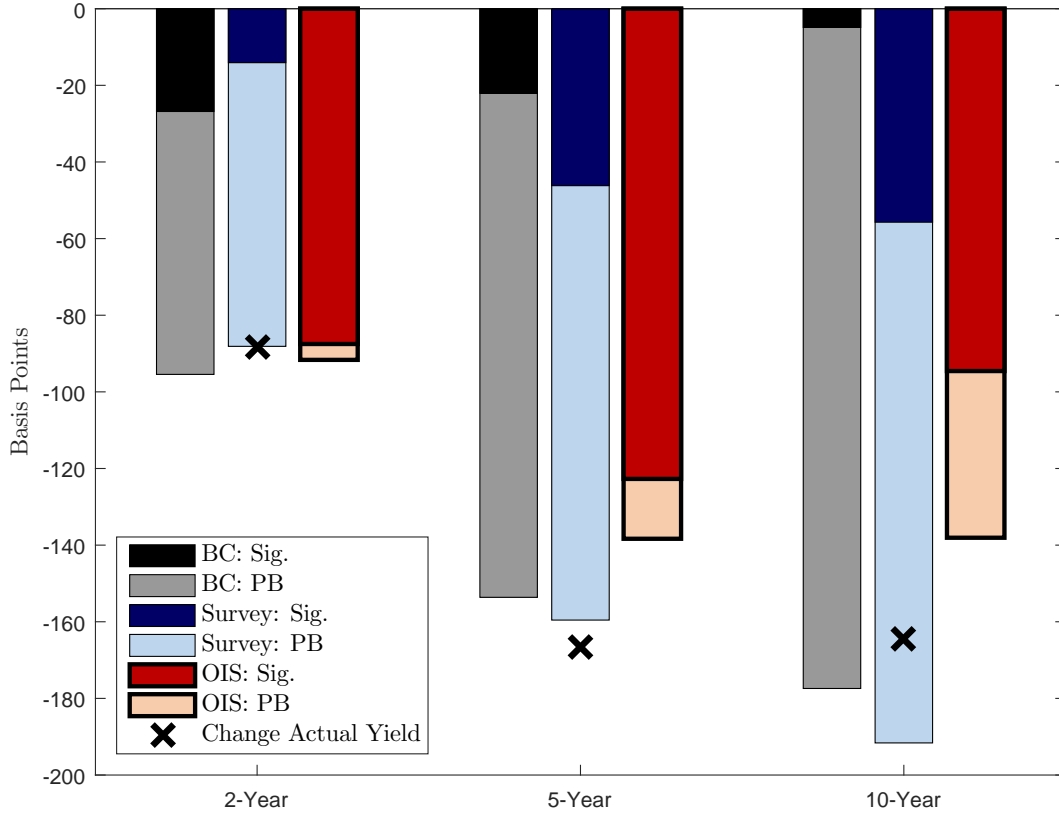
10-year OIS rates declined by 152.05 and 162.15 basis points in turn.

Tables 4.4 also provides some indication of the changing efficacy of unconventional monetary policy announcements over time. In particular, actual Treasury yields fell most in response to earlier policy announcements, linked with QE1. The largest fall in Treasury yields came on event day V, when the Fed announced the extension of QE1 to include the purchase of long-term Treasuries, and altered their forward guidance from stating that the federal funds rate would remain low, from “for some time” to “an extended period.” The introduction of calendar-based forward guidance on event day XI was the only to depress bond yields by a similar order of magnitude to the early announcements.

4.4.2 Event Study Results

To explicitly consider the reaction of interest rate expectations and term premia to policy announcements, I perform an event study using the three GADTSM decompositions described in section 4.3. As in the existing literature, the risk-neutral yields are associated with the signalling channel and term premia with portfolio rebalancing. The results are reported in tables

Figure 4.2: Cumulative Fall in Yields on 11 Event Days with Statistically Significant Falls in Fitted Yields



Note: Plot of the cumulative fall in actual yields, fitted yields, risk-neutral yields and term premia on 11 event days with statistically significant falls in fitted yields. The 11 dates exclude events IV, VII, IX, XII and XVI. The remaining 11 event days are listed in table 4.3. The cumulative fall in actual 2, 5 and 10-year yields is denoted by a black cross. The cumulative fall in fitted 2, 5 and 10-year yields from the bias-corrected ('BC') (Bauer et al., 2012), survey-augmented ('Survey') (Kim and Wright, 2005) and OIS-augmented ('OIS') (Lloyd, 2017a) GADTSMs are depicted in the bars. The upper, dark, segment of each bar depicts the cumulative fall in risk-neutral yields — the signalling channel ('Sig.'). The lower, light, segment of each bar depicts the cumulative fall in term premia — the portfolio balance channel ('PB').

4.5, 4.6 and 4.7 for the 2, 5 and 10-year yields respectively. To assess the statistical significance of daily changes, I control for major US macroeconomic data release dates.²⁹ All three tables indicate that changes in fitted yields and their sub-components on event days were sizeable, and statistically significant on many event dates. Notably, both the signalling and portfolio balance channels are shown to be operative, albeit to differing degrees at different horizons and with different yield curve decompositions.

Figure 4.2 graphically depicts the headline results. It plots the cumulative fall in actual 2, 5 and 10-year yields (denoted by a cross) along with the cumulative fall in fitted yields, risk-neutral yields and term premia from the three term structure models on the 11 event days when bond yields fell significantly. These 11 dates exclude events IV, VII, IX, XII and XVI, when fitted yields either increased or fell insignificantly. The graph demonstrates that the survey-augmented decomposition attributes a greater proportion of variation in fitted yields to the term premium, and thus portfolio rebalancing. It attributes 84.01% of the cumulative fall in the 2-year fitted yield on the 11 event days to the term premium, the highest proportion of the three models at this maturity. The corresponding figures for the 5 and 10-year horizons are 71.05%

²⁹See appendix C.1.2 for further details.

and 70.91% respectively. However, for the reasons outlined in section 4.3, the results from the survey-augmented decomposition over-attribute variation in fitted yields to the term premium, falsely overstating the efficacy of the portfolio balance channel relative to signalling.

Although the bias-corrected decomposition attributes a lesser percentage of fitted yield variation to term premia at all three horizons than the survey-augmented decomposition on all 16 event dates, it attributes the greatest proportion of variation in fitted yields to the term premium at the 5 and 10-year horizons on the 11 dates when fitted yields fell significantly. At the 2, 5 and 10-year horizons the bias-corrected decomposition respectively attributes 71.89%, 85.63% and 97.24% of the cumulative fall in fitted yields term premia, and thus portfolio rebalancing, on these 11 days.

The preferred OIS-augmented decomposition highlights a powerful role for signalling at all horizons. Of the cumulative fall in fitted yields at the 2, 5 and 10-year horizons on the 11 event days with significant falls in fitted yields, the OIS-augmented decomposition respectively attributes 95.49%, 88.71% and 68.49% to falls in risk-neutral yields and thus signalling. The corresponding figures for the complete set of 16 event dates are 95.29%, 90.61% and 74.19%. The OIS-augmented model attributes the greatest proportion of variation to signalling at the 2-year horizon. But in terms of absolute size, the reaction of risk-neutral yields is hump-shaped with respect to bond maturity. The 2-year risk-neutral yield fell by a total of 87.51 basis points on the 11 event days, while the 5 and 10-year figures were 122.72 and 94.57 basis points respectively, indicating that interest rate expectations were affected well beyond the 2-year horizon. The relative and absolute size of term premium changes are increasing in maturity, reflecting greater importance of risk at longer horizons.

Tables 4.5-4.7 provide a more detailed breakdown of the event study, allowing a comparison of LSAP and forward guidance event days. For all three maturities, falls in risk-neutral yields from the OIS-augmented decomposition explain a marginally larger proportion of falls in fitted yields on forward guidance event days than LSAP event days. For instance, falls in risk-neutral yields explain 91.31% of the reduction in the 5-year fitted yield on forward guidance days, and 89.96% on LSAP days. However, this comparison is blurred because most forward guidance events also included some information about LSAPs. Nevertheless, it is particularly striking that proportional expectations effects were strong on LSAP-only event days (events I, II, VI, VIII and X especially). For instance, on event date VIII, when the Fed announced it would reinvest maturing assets to maintain the stock of asset purchases, over 60% of the fall in the 10-year fitted yield is attributed to a reduction in the risk-neutral yield.

Tables 4.5-4.7 indicate some interesting differences between each of the three yield curve decompositions on specific event days too. For example, on event day II (December 1, 2008), Fed Chairman Ben Bernanke stated that “US Treasuries may be purchased” as part of the LSAP program. The survey-augmented decomposition attributes just 15% of the fall in the 2-year fitted yield on that day to the risk-neutral yield, whereas the OIS-augmented model attributes around 94% to the fall in the risk-neutral yield. According to this latter decomposition, the signalling effect of this announcement was more pronounced. Similar differences exist on event day V (March 18, 2009), when the Fed stated that low interest rates would likely be warranted for “an extended period” and announced the purchase of longer-term Treasuries as part of the LSAP program. On this day, the bias-corrected and survey-augmented models respectively attribute 19% and 16% of the fall in the 2-year fitted yield to the risk-neutral yield. In contrast, the

Table 4.5: 2-Year US Treasury Yield and its Components: Changes on Event Dates for Three Decompositions of US Zero-Coupon Treasury Yields (All Figures in Basis Points to 2 Decimal Places)

#	Event Date & Type	GADTSM Decomposition									
		Bias-Corrected			Survey-Augmented			OIS-Augmented			
		(1) $\widehat{\Delta yld_2}$	(2) $\widehat{\Delta exp_2}$	(3) $\widehat{\Delta \hat{tp}_2}$	(4) $\widehat{\Delta yld_2}$	(5) $\widehat{\Delta exp_2}$	(6) $\widehat{\Delta \hat{tp}_2}$	(7) $\widehat{\Delta yld_2}$	(8) $\widehat{\Delta exp_2}$	(9) $\widehat{\Delta \hat{tp}_2}$	
I	25/11/2008	L	-14.63***	-7.79***	-6.84***	-14.35***	-4.11***	-10.24***	-12.23***	-11.88***	-0.35***
II	01/12/2008	L	-12.98***	-7.05***	-5.93***	-11.86***	-1.78***	-10.08***	-12.81***	-12.00***	-0.81***
III	16/12/2008	F,L	-11.55***	-6.35***	-5.19***	-10.72***	-4.17***	-6.55***	-11.15***	-11.04***	-0.11***
IV	28/01/2009	L	3.68***	-6.12***	9.79***	4.57***	0.17**	4.40***	5.31***	5.19***	0.12***
V	18/03/2009	F,L	-28.08***	-5.34**	-22.74***	-26.41***	-4.29**	-22.12***	-26.30***	-25.35***	-0.95***
VI	10/08/2010	L	-3.09**	0.95	-4.05***	-2.68*	0.00	-2.68***	-2.86**	-2.71**	-0.15***
VII	28/08/2010	L	5.72***	-3.82	9.54***	5.39***	-0.65	6.04***	6.35***	5.97***	0.37***
VIII	21/09/2010	L	-4.03***	2.59***	-6.62***	-3.71***	0.43***	-4.14***	-4.67***	-4.33***	-0.34***
IX	15/10/2010	F,L	-1.05*	-9.89***	8.84***	-1.21**	-2.59***	1.38***	-0.01	0.12	-0.12***
X	03/11/2010	L	-2.58***	-6.89***	4.31***	-1.51***	-0.97***	-0.54***	-1.93***	-1.63***	-0.29***
XI	09/08/2011	F,L	-9.26***	2.46*	-11.72***	-8.56***	0.09	-8.65***	-9.32***	-8.92***	-0.40***
XII	26/08/2011	F,L	-1.64	0.60**	-2.24**	-1.70	0.22**	-1.92	-2.65	-2.42	-0.23***
XIII	25/01/2012	F	-4.18***	0.26	-4.45***	-3.78***	0.24***	-4.02***	-4.74***	-4.34***	-0.40***
XIV	31/08/2012	F,L	-3.70***	-0.45***	-3.25***	-3.67***	-0.50***	-3.17***	-3.72***	-3.58***	-0.14***
XV	13/09/2012	F,L	-1.35*	0.78	-2.13***	-0.87	0.97***	-1.84***	-1.93***	-1.73***	-0.20***
XVI	12/12/2012	F	-0.15	-4.75***	4.60***	0.03	-1.58***	1.61***	0.23*	0.12	0.11***
Total Change on Event Days											
LSAP Events			-84.55	-46.32 (54.78%)	-38.24 (45.22%)	-77.29	-17.18 (22.23%)	-60.11 (77.78%)	-77.91	-74.32 (95.39%)	-3.60 (4.61%)
FG Events			-60.96	-22.68 (37.21%)	-38.28 (62.79%)	-56.89	-11.61 (20.41%)	-45.28 (79.59%)	-59.59	-57.16 (95.91%)	-2.43 (4.09%)
All Events			-88.89	-50.80 (57.16%)	-38.08 (42.84%)	-81.04	-18.52 (22.85%)	-62.52 (77.15%)	-82.43	-78.54 (95.29%)	-3.88 (4.71%)
Proportion of total change in fitted yield $\widehat{\Delta yld_2}$ explained by $\widehat{\Delta exp_2}$ and $\widehat{\Delta \hat{tp}_2}$ shown in brackets beneath.											

Proportion of total change in fitted yield $\Delta \widehat{yld}_2$ explained by $\Delta \widehat{exp}_2$ and $\Delta \widehat{tp}_2$ shown in brackets beneath.

Abbreviations: L = Event date with LSAP news; F = Event date with forward guidance announcements.

$\Delta \widehat{yld}_2$: Change in 2-year fitted yield on event dates from the bias-corrected Bauer et al. (2012), survey-augmented Kim and Wright (2005), and OIS-augmented Lloyd (2017a) decompositions. $\Delta \widehat{exp}_2$ ($\Delta \widehat{tp}_2$): Change in 2-year risk-neutral yield (term premium) on event days. Tests to determine the significance of daily changes are described in appendix C.1.2; t -statistics are calculated using Newey and West (1987) standard errors. Daily changes that are significant at the 1%, 5% and 10% levels are denoted with asterisks ***, ** and * respectively.

Table 4.6: 5-Year US Treasury Yield and its Components: Changes on Event Dates for Three Decompositions of US Zero-Coupon Treasury Yields (All Figures in Basis Points to 2 Decimal Places)

#	Event Date	Bias-Corrected			GADTSM Decomposition Survey-Augmented			OIS-Augmented			
	$\&$ Type	(1) $\widehat{\Delta yld_5}$	(2) $\Delta \widehat{exp_5}$	(3) $\Delta \widehat{tp_5}$	(4) $\widehat{\Delta yld_5}$	(5) $\Delta \widehat{exp_5}$	(6) $\Delta \widehat{tp_5}$	(7) $\widehat{\Delta yld_5}$	(8) $\Delta \widehat{exp_5}$	(9) $\Delta \widehat{tp_5}$	
I	25/11/2008	L	-21.33***	-8.03***	-13.30***	-21.57***	-6.77***	-14.80***	-17.66***	-15.90***	-1.76***
II	01/12/2008	L	-21.14***	-8.31***	-12.83***	-20.53***	-5.57***	-14.96***	-19.07***	-17.20***	-1.87***
III	16/12/2008	F,L	-16.42***	-6.47***	-9.95***	-16.16***	-6.15***	-10.01***	-15.06***	-13.91***	-1.15***
IV	28/01/2009	L	8.00***	-9.16***	17.16***	9.43***	2.51***	6.92***	7.57***	6.84***	0.74***
V	18/03/2009	F,L	-45.23***	-2.61	-42.62***	-47.07***	-13.48***	-33.59***	-39.03***	-34.77***	-4.26***
VI	10/08/2010	L	-5.98***	1.63	-7.61***	-6.52***	-1.94***	-4.58***	-5.44***	-4.32***	-1.12***
VII	28/08/2010	L	11.98***	-5.75	17.73***	12.63***	2.99***	9.64***	10.11***	8.78***	1.33***
VIII	21/09/2010	L	-8.39***	3.92***	-12.31***	-9.12***	-2.31***	-6.81***	-8.00***	-6.74***	-1.26***
IX	15/10/2010	F,L	2.58***	-13.03***	15.62***	3.68***	0.38	3.30***	2.42***	0.93	1.49***
X	03/11/2010	L	-1.67***	-8.93***	7.25***	-1.92***	-1.04***	-0.88***	-2.43***	-2.52***	0.09***
XI	09/08/2011	F,L	-16.91***	4.71***	-21.63***	-18.43***	-4.73***	-13.70***	-16.50***	-13.50***	-3.01***
XII	26/08/2011	F,L	-3.19	1.00**	-4.19**	-3.50	-0.66	-2.84	-3.62	-3.46	-0.16
XIII	25/01/2012	F	-7.69***	0.81**	-8.50***	-8.42***	-2.06***	-6.36***	-7.23***	-6.51***	-0.72***
XIV	31/08/2012	F,L	-6.07***	0.00	-6.07***	-6.36***	-1.68***	-4.68***	-5.71***	-5.00***	-0.71***
XV	13/09/2012	F,L	-2.80***	1.19	-3.98***	-3.46***	-0.46	-3.00***	-2.21***	-2.36***	0.15
XVI	12/12/2012	F	2.56***	-6.00***	8.56***	2.89***	0.09*	2.80***	2.12***	1.13***	0.99***
Total Change on Event Days											
LSAP Events		-126.56	-49.84	-76.72	-128.90	-38.91	-89.99	-114.63	-103.12	-11.51	
			(39.38%)	(60.62%)		(30.19%)	(69.81%)		(89.96%)	(10.04%)	
FG Events		-93.17	-20.41	-72.77	-96.83	-28.75	-68.08	-84.83	-77.45	-7.38	
			(21.90%)	(78.10%)		(29.69%)	(70.31%)		(91.31%)	(8.69%)	
All Events		-131.69	-55.03	-76.66	-134.43	-40.88	-93.55	-119.74	-108.50	-11.24	
			(41.79%)	(58.21%)		(30.41%)	(69.59%)		(90.61%)	(9.39%)	
Proportion of total change in fitted yield $\widehat{\Delta yld_5}$ explained by $\Delta \widehat{exp_5}$ and $\Delta \widehat{tp_5}$ shown in brackets beneath.											
Abbreviations: L = Event date with LSAP news; F = Event date with forward guidance announcements.											
$\widehat{\Delta yld_5}$: Change in 5-year fitted yield on event dates from the bias-corrected Bauer et al. (2012), survey-augmented Kim and Wright (2005), and OIS-augmented Lloyd (2017a) decompositions. $\Delta \widehat{exp_5}$ ($\Delta \widehat{tp_5}$): Change in 5-year risk-neutral yield (term premium) on event days. Tests to determine the significance of daily changes are described in appendix C.1.2; t -statistics are calculated using Newey and West (1987) standard errors. Daily changes that are significant at the 1%, 5% and 10% levels are denoted with asterisks ***, ** and * respectively.											

Table 4.7: 10-Year US Treasury Yield and its Components: Changes on Event Dates for Three Decompositions of US Zero-Coupon Treasury Yields (All Figures in Basis Points to 2 Decimal Places)

#	Event Date & Type	GADTSM Decomposition									
		Bias-Corrected			Survey-Augmented			OIS-Augmented			
		(1) $\widehat{\Delta yld}_{10}$	(2) $\Delta \widehat{exp}_{10}$	(3) $\Delta \widehat{fp}_{10}$	(4) $\widehat{\Delta yld}_{10}$	(5) $\Delta \widehat{exp}_{10}$	(6) $\Delta \widehat{fp}_{10}$	(7) $\widehat{\Delta yld}_{10}$	(8) $\Delta \widehat{exp}_{10}$	(9) $\Delta \widehat{fp}_{10}$	
I	25/11/2008	L	-22.62***	-6.56	-16.06***	-23.91***	-7.17***	-16.74***	-17.43***	-12.14***	-5.29***
II	01/12/2008	L	-21.79***	-7.22***	-14.57***	-24.01***	-6.58***	-17.43***	-17.64***	-13.30***	-4.34***
III	16/12/2008	F,L	-17.36***	-5.35***	-12.00***	-18.44***	-6.54***	-11.90***	-14.32***	-10.48***	-3.84***
IV	28/01/2009	L	13.89***	-12.26***	26.15***	11.82***	3.30***	8.52***	7.50***	5.20***	2.30***
V	18/03/2009	F,L	-54.14***	3.37	-57.51***	-56.06***	-16.11***	-39.95***	-39.06***	-26.68***	-12.38***
VI	10/08/2010	L	-7.93***	2.77	-10.70***	-8.70***	-2.68***	-6.02***	-6.89***	-3.41***	-3.48***
VII	28/08/2010	L	16.97***	-8.53***	25.50***	16.30***	4.29***	12.01***	10.43***	6.83***	3.60***
VIII	21/09/2010	L	-11.84***	5.85***	-17.69***	-12.03***	-3.35***	-8.68***	-8.74***	-5.30***	-3.44***
IX	15/10/2010	F,L	8.73***	-16.00***	24.73***	7.09***	1.80***	5.29***	6.32***	0.85	5.47***
X	03/11/2010	L	1.85***	-10.51***	12.36***	-2.11***	-1.01***	-1.10***	-0.84***	-1.97***	1.12***
XI	09/08/2011	F,L	-22.65***	8.06***	-30.71***	-23.34***	-6.41***	-16.93***	-19.92***	-10.56***	-9.36***
XII	26/08/2011	F,L	-4.30	1.63**	-5.93***	-4.25	-0.93	-3.32*	-2.62	-2.68	0.06
XIII	25/01/2012	F	-9.63***	2.01***	-11.65***	-10.74***	-2.87***	-7.87***	-6.48***	-5.08***	-1.41***
XIV	31/08/2012	F,L	-7.42***	0.87*	-8.29***	-7.42***	-1.98***	-5.44***	-5.98***	-3.85***	-2.13***
XV	13/09/2012	F,L	-3.89***	1.81	-5.69***	-4.86***	-1.03***	-3.83***	-0.76*	-1.80***	1.04**
XVI	12/12/2012	F	5.54***	-7.47***	13.01***	4.57***	0.82***	3.75***	4.11***	1.02***	3.08***
Total Change on Event Days											
LSAP Events			-132.48	-42.06	-90.42	-149.92	-44.40	-105.52	-109.95	-79.28	-30.67
				(31.75%)	(68.25%)		(29.62%)	(70.38%)		(72.11%)	(27.89%)
FG Events			-105.10	-11.06	-94.05	-113.45	-33.25	-80.20	-78.71	-59.26	-19.45
				(10.52%)	(89.48%)		(29.31%)	(70.69%)		(75.28%)	(24.72%)
All Events			-136.57	-47.51	-89.07	-156.09	-46.45	-109.64	-112.33	-83.34	-28.99
				(34.79%)	(65.21%)		(29.76%)	(70.24%)		(74.19%)	(25.81%)
Proportion of total change in fitted yield Δyld_{10} explained by $\Delta \widehat{exp}_{10}$ and $\Delta \widehat{fp}_{10}$ shown in brackets beneath.											

Proportion of total change in fitted yield $\Delta \widehat{yld}_{10}$ explained by $\Delta \widehat{exp}_{10}$ and $\Delta \widehat{fp}_{10}$ shown in brackets beneath.

Abbreviations: L = Event date with LSAP news; F = Event date with forward guidance announcements.

$\Delta \widehat{yld}_{10}$: Change in 10-year fitted yield on event dates from the bias-corrected Bauer et al. (2012), survey-augmented Kim and Wright (2005), and OIS-augmented Lloyd (2017a) decompositions. $\Delta \widehat{exp}_{10}$ ($\Delta \widehat{fp}_{10}$): Change in 10-year risk-neutral yield (term premium) on event days. Tests to determine the significance of daily changes are described in appendix C.1.2; t -statistics are calculated using Newey and West (1987) standard errors. Daily changes that are significant at the 1%, 5% and 10% levels are denoted with asterisks ***, ** and * respectively.

corresponding figure for the OIS-augmented decomposition is 96%.

In sum, the event study evidence indicates that different yield curve decompositions provide differing conclusions about the relative efficacy of signalling and portfolio rebalancing. Using the preferred OIS-augmented model, I find that unconventional monetary policy announcements had particularly powerful signalling effects on financial markets, explaining between 68.49% and 95.49% of the cumulative decline in bond yields on announcement days.

4.5 Signalling, Portfolio Rebalancing and the Real Economy

To assess the relative importance of signalling and portfolio rebalancing effects of unconventional monetary policy for the real economy, I identify the shocks with a combination of zero-impact and sign restrictions within a structural vector autoregression (SVAR).³⁰

4.5.1 SVAR Methodology

The baseline reduced-form VAR, labelled **model 1**, consists of four monthly time series:

$$\mathbf{Y}_t = [ip_t, p_t, exp_t^{(j)}, tp_t^{(j)}]' \quad (4.2)$$

where ip_t is the logarithm of industrial production and p_t is the logarithm of the consumer price index (CPI).³¹ Because the federal funds futures rate reached its ELB in December 2008, I do not include this in the set of variables. Instead, I use the components of longer-term interest rates to indicate the stance of monetary policy over the sample period. $exp_t^{(j)}$ represents the monthly average of the j -year risk-neutral yield estimated from one of the three GADTSMs described in section 4.3, while $tp_t^{(j)}$ denotes the monthly average of the j -year term premium from the same GADTSM. In accordance with the existing unconventional monetary policy VAR literature (e.g. Baumeister and Benati, 2013; IMF, 2013; Lloyd, 2013), I use the 10-year US Treasury yield in my baseline analysis. To assess the robustness of my results, I also run the SVAR with decompositions of different horizon yields.

I estimate the VAR using data from November 2008 to April 2013, the period in which LSAP and forward guidance announcements occurred and policy rates were at the ELB.³² In line with the Schwarz-Bayes information criterion, the lag order of the VAR is two.

To further assess the robustness of my results, I account for additional channels through which LSAPs and forward guidance may transmit to the real economy. I control for each additional channel in turn.

I account for the credit channel by extending the baseline VAR (4.2) to form a five-variable system, labelled **model 2**:

$$\mathbf{Y}_t = [ip_t, p_t, exp_t^{(j)}, tp_t^{(j)}, cred_t]' \quad (4.3)$$

³⁰To impose these restrictions, I use the algorithm of Binning (2013). Arias, Rubio-Ramírez, and Waggoner (2014) state that this algorithm for zero-impact and sign restrictions does not impose extra hidden sign restrictions on the model, unlike other existing approaches.

³¹Data sources are provided in appendix A.1. Variables are included in (log) levels, consistent with Sims, Stock, and Watson (1990) who show that parameter estimates from a VAR with potentially non-stationary log-level variables (e.g. ip_t and p_t) are consistent.

³²The November 2008 sample start date is defined by the first LSAP announcement in table 4.3. The April 2013 sample end date is chosen because of the May 2013 ‘taper tantrum’. Nevertheless, the results are robust when December 2015 is chosen as the end date.

where $cred_t$ is the logarithm of US bank credit, a measure of bank lending.

Model 3 accounts for international effects. Bauer and Neely (2012) argue that the exchange rate channel of unconventional monetary policy works through international interest rate differentials, so is a component of signalling and portfolio balance channels. For this reason, I omit real exchange rate from models 1 and 2. However, in model 3, I assess the robustness of my findings to the inclusion of international factors:

$$\mathbf{Y}_t = [ip_t, p_t, exp_t^{(j)}, tp_t^{(j)}, rer_t]' \quad (4.4)$$

where rer_t represents the logarithm of the effective exchange rate series for the US against 60 other countries.³³

4.5.2 Sign Restrictions

For the baseline specification (4.2), I identify four structural shocks: aggregate demand; aggregate supply; signalling; and portfolio rebalancing. It is important that these structural shocks are separately identified, posing a challenge for the identification of signalling and portfolio balance shocks. Given this, I assess the robustness of my results to two identification schemes, summarised in table 4.8.

Table 4.8: Sign Restriction Schemes 1 & 2 for the Four Variable VAR (4.2) - Model 1

Shock	Variables					
	Schemes 1 & 2		Scheme 1		Scheme 2	
	ip	p	$exp^{(j)}$	$tp^{(j)}$	$exp^{(j)}$	$tp^{(j)}$
Demand	> 0	> 0	> 0	\cdot	> 0	\cdot
Supply	> 0	< 0	\cdot	\cdot	\cdot	\cdot
Signalling	0	0	> 0	\cdot	> 0	$\mathbf{0}$
Portfolio Balance	0	0	$\mathbf{0}$	> 0	\cdot	> 0

\cdot denotes an unrestricted response. 0 denotes a response that is restricted to zero in the month of the shock. $< (>)$ denotes a response that is strictly negative (positive) in the month of the shock.

To identify the first two structural shocks, to aggregate demand and supply, I appeal to a standard set of sign restrictions used in a number of studies of conventional and unconventional monetary policy (e.g. Baumeister and Benati, 2013). For these shocks, I use the same identification restrictions in schemes 1 and 2. A positive demand shock must increase industrial production and the price level. Moreover, since the shock should tighten future monetary policy (conventional or unconventional), I impose that the expected future path of short-term interest rates $exp^{(j)}$ must also increase. A supply shock must have opposing effects on industrial production and the price level, while the response of all other variables is left unrestricted.

Signalling and portfolio balance shocks emanate from changes in the expected future short-term interest rate component $exp^{(j)}$ and the term premium $tp^{(j)}$ respectively. Section 4.4.2 showed that unconventional monetary policy announcements did significantly reduce both components of longer-term yields, ratifying the assumption that shocks to these components in the

³³An increase in rer_t corresponds to an appreciation of the US real exchange rate.

VAR can be associated with policy. The effects of these two shocks on industrial production and the price level are the same for identification schemes 1 and 2. Contractionary signalling or portfolio balance shocks are constrained to have no effect on industrial production or the price level on impact to ensure that the time path of their reaction to shocks is realistic. Thereafter, the sign of the shock is unconstrained. This is a useful feature, because the sign and size of the responses of industrial production and the price level to signalling and portfolio rebalancing shocks will be used as metrics to gauge the relative efficacy of the two sub-channels.

To separately identify the signalling and portfolio balance shocks, the $exp^{(j)}$ and $tp^{(j)}$ columns of table 4.8 must differ. In identification scheme 1, I impose that the portfolio balance shock has no instantaneous impact on the expected future path of short-term interest rates. This is labelled a ‘pure term premium shock’, such that shocks to the term premia on a j -period bond do not alter investors’ expectations of future short-term interest rates over the same horizon with a given month. To assess the sensitivity of my results to this restriction, I impose its converse in scheme 2: signalling shocks cannot exert a contemporaneous impact on the term premium. This is labelled a ‘pure expectations shock’. Of the two schemes, scheme 1 is most plausible. The term premium includes interest rate risk, which is likely to be immediately influenced by the signalling shock. However, to the extent that both expectations and term premia significantly responded to unconventional monetary policy announcements, scheme 1 may provide a more realistic quantitative assessment of signalling shocks, while scheme 2 may be better suited to capturing portfolio rebalancing shocks.

When controlling for bank credit and the real exchange rate in models 2 and 3, I do not identify additional shocks. In model 2, I restrict the impact response of bank credit to signalling and portfolio rebalancing shocks to zero, motivated by evidence that bank lending reacts to monetary policy with a lag (Bernanke and Blinder, 1992). The response of bank credit in subsequent periods is unrestricted in size and sign. For model 3, I impose that the real exchange rate must depreciate on impact in response to expansionary signalling and portfolio rebalancing shocks, but leave the response thereafter unrestricted. This is motivated by evidence that nominal exchange rates depreciated in response to US LSAP surprises (Glick and Leduc, 2012), implying real exchange rate depreciation in the presence of price stickiness.

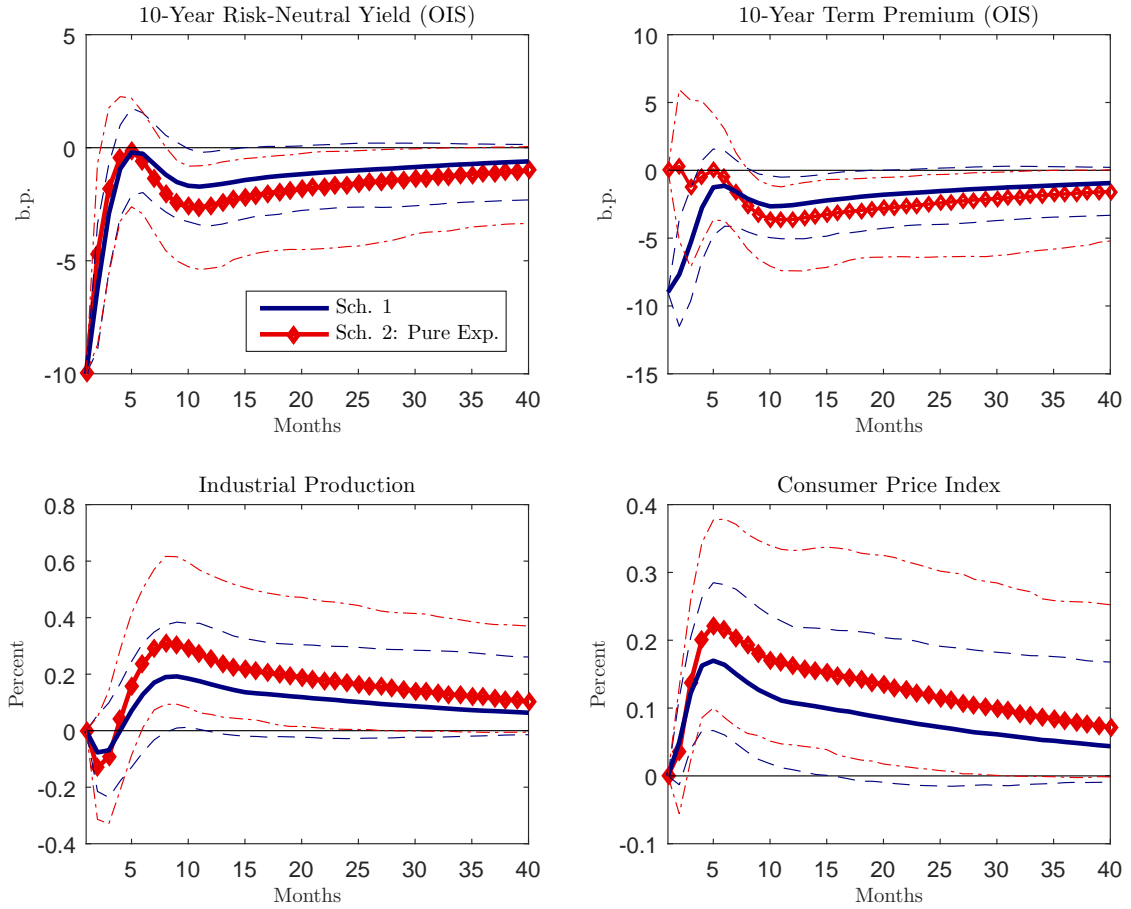
4.5.3 SVAR Results

The sign-restricted SVAR results show that, contrary to the prior beliefs of policymakers (Bernanke, 2010), reductions in longer-term interest rates between November 2008 and April 2013 have exerted a more powerful effect on US industrial production via the signalling channel.

Model 1: 4 Variable SVAR

Figures 4.3 and 4.4 present the median impulse response functions, together with the 5% and 95% confidence intervals, for the signalling and portfolio rebalancing shocks from model 1, estimated using the preferred OIS-augmented 10-year yield decomposition. To foster comparison, the median shocks are normalised to represent a 10 basis point fall in the corresponding yield

Figure 4.3: Impulse Response Functions to a Signalling Shock for Model 1 using Sign Restriction Schemes 1 & 2 and the OIS-Augmented Decomposition of the 10-Year Treasury Yield



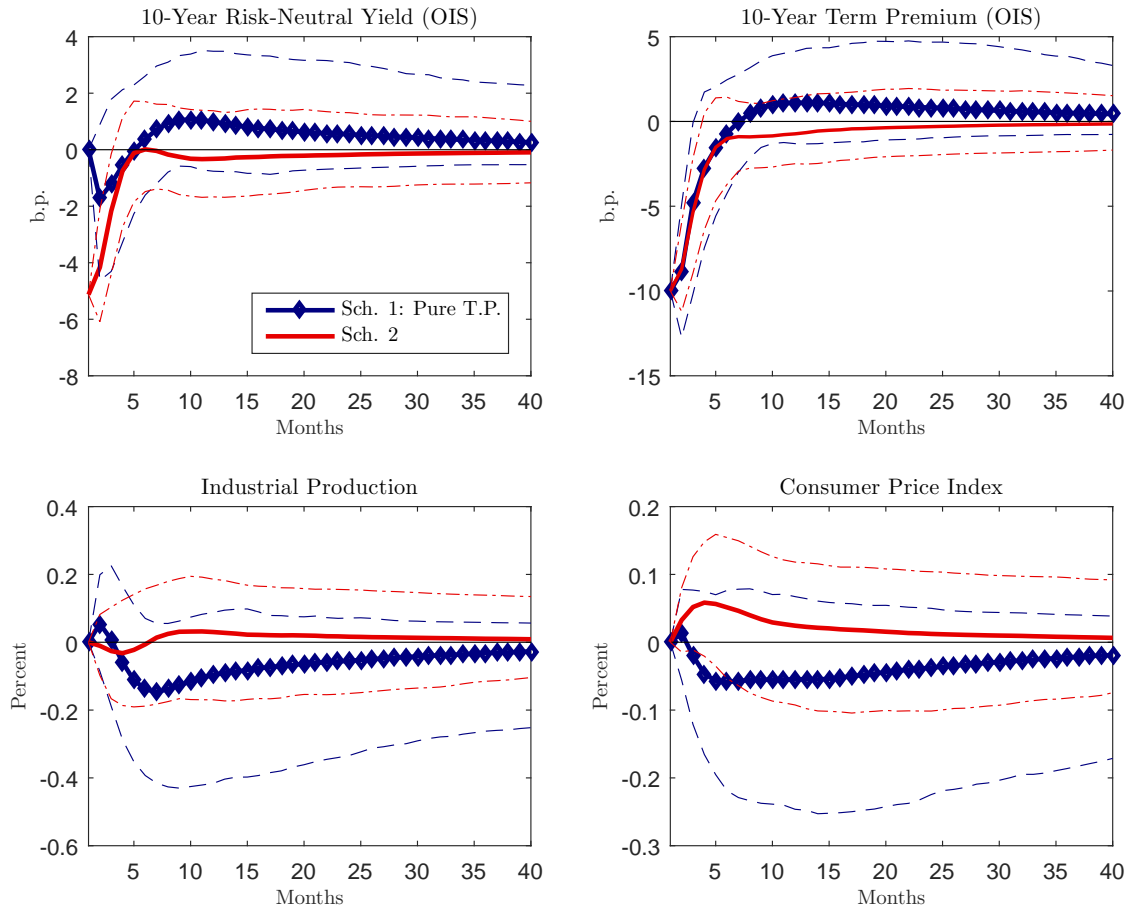
Note: The plots show the impulse response functions to a signalling shock, normalised to represent a 10 basis point fall in the 10-year risk-neutral yield, estimated using the OIS-augmented decomposition (Lloyd, 2017a). The VAR is estimated using data from November 2008 to April 2013. The two sign restriction schemes are detailed in table 4.8. The bold lines denote median impulse response draws. The thin dashed lines represent the 5% and 95% confidence intervals around median impulse responses, constructed using a residual-based block bootstrap.

component.³⁴ Within the VAR, the average proportion of historical variation in industrial production attributed to the two long-term interest rate shocks is 29.5% and 5.4% under schemes 1 and 2 respectively.

Figure 4.3 documents that the signalling shock has a significant lagged positive effect on industrial production and CPI using both identification schemes. With scheme 1, the response of industrial production is significantly positive 9 to 11 months after the shock, peaking at 0.19% after 9 months. The peak response of industrial production under scheme 2, the pure expectations shock, is larger at 0.30% after 8 months, and the impulse response is significantly positive 6 to 27 months after the shock. The CPI also significantly increases 3 to 15 months after the shock under scheme 1, and 3 to 31 months under scheme 2. Its peak responses, 5 months after the shock, are 0.17% and 0.23% under schemes 1 and 2 respectively.

³⁴The median response of the expectations (term premium) component for a signalling (portfolio rebalancing) shock is similar in quantity. The two responses have been equalised to enable discussion of the effects of equal-sized shocks. I use a residual-based block bootstrap with 1000 replications and 100 rotations per bootstrap, which Brüggemann, Jentsch, and Trenkler (2016) demonstrate leads to asymptotically valid inference on structural impulse response functions in the presence of conditional heteroskedasticity, unlike a wild or pairwise bootstrap.

Figure 4.4: Impulse Response Functions to a Portfolio Rebalancing Shock for Model 1 using Sign Restriction Schemes 1 & 2 and the OIS-Augmented Decomposition of the 10-Year Treasury Yield



Note: The plots show the impulse response functions to a portfolio rebalancing shock, normalised to represent a 10 basis point fall in the 10-year term premium, estimated using the OIS-augmented decomposition (Lloyd, 2017a). The VAR is estimated using data from November 2008 to April 2013. The two sign restriction schemes are detailed in table 4.8. The bold lines denote median impulse response draws. The thin dashed lines represent the 5% and 95% confidence intervals around median impulse responses, constructed using a residual-based block bootstrap.

In contrast to the signalling shock, the responses of industrial production and CPI to an expansionary portfolio rebalancing shock in figure 4.4 are statistically insignificant at all horizons. Moreover, their peak responses are considerably smaller than their responses to the signalling shock. Under scheme 1, the pure term premium shock, the peak responses of industrial production and CPI are 0.09% and 0.01% respectively. The corresponding figures for scheme 2 are 0.03% and 0.06%. All are insignificantly different from zero. Comparing the peak responses of industrial production to the two shocks indicates that the signalling shock explains 68% to 91% of the total peak industrial production effect due to the long-term interest rate shocks.

The signalling shock also accounts for a greater proportion of industrial production forecast error variation than the portfolio rebalancing shock in the year following a shock under both identification schemes. With scheme 1, the signalling shock explains 77.8% of industrial production forecast error variation attributed to the two long-term interest rate shocks after 3 months. Although the peak figure under scheme 2 is smaller, the signalling shock, on average, explains 58.4% of industrial production forecast error variation due to the long-term interest rate shocks in the 12-month period after the perturbation. Thus, the signalling channel is associated with

around two-thirds to three-quarters of the total effects of long-term interest rate shocks on US industrial production between November 2008 and April 2013.

In summary, when model 1 is estimated using the OIS-augmented decomposition of the 10-year Treasury yield, the signalling channel is associated with the majority of total industrial production effects attributed to shocks to longer-term interest rates. Lloyd (2017c, Appendix C.1) presents robustness exercises. It illustrates that the results strengthen when a longer sample period (November 2008 to December 2015) is used. Extending the sample period makes the effects of the signalling shock on industrial production and CPI stronger and significant for a longer horizon. The results are also robust to the use of different maturity yields and different yield curve decompositions. Even for the 2-year survey-augmented model, which attributes a smaller fraction of variation in interest rates to signalling in the event study, the signalling shock has stronger effects on industrial production and CPI than the portfolio rebalancing shock. The results are also robust to using the 2-year OIS rate, instead of the 2-year risk-neutral yield, alongside the 2-year term premium in the VAR. This robustness check accounts for the possibility that, because both the risk-neutral yield and the term premium are estimated within the same GADSTM, they do not vary independently.

4.5.4 Models 2 and 3: Adding Controls

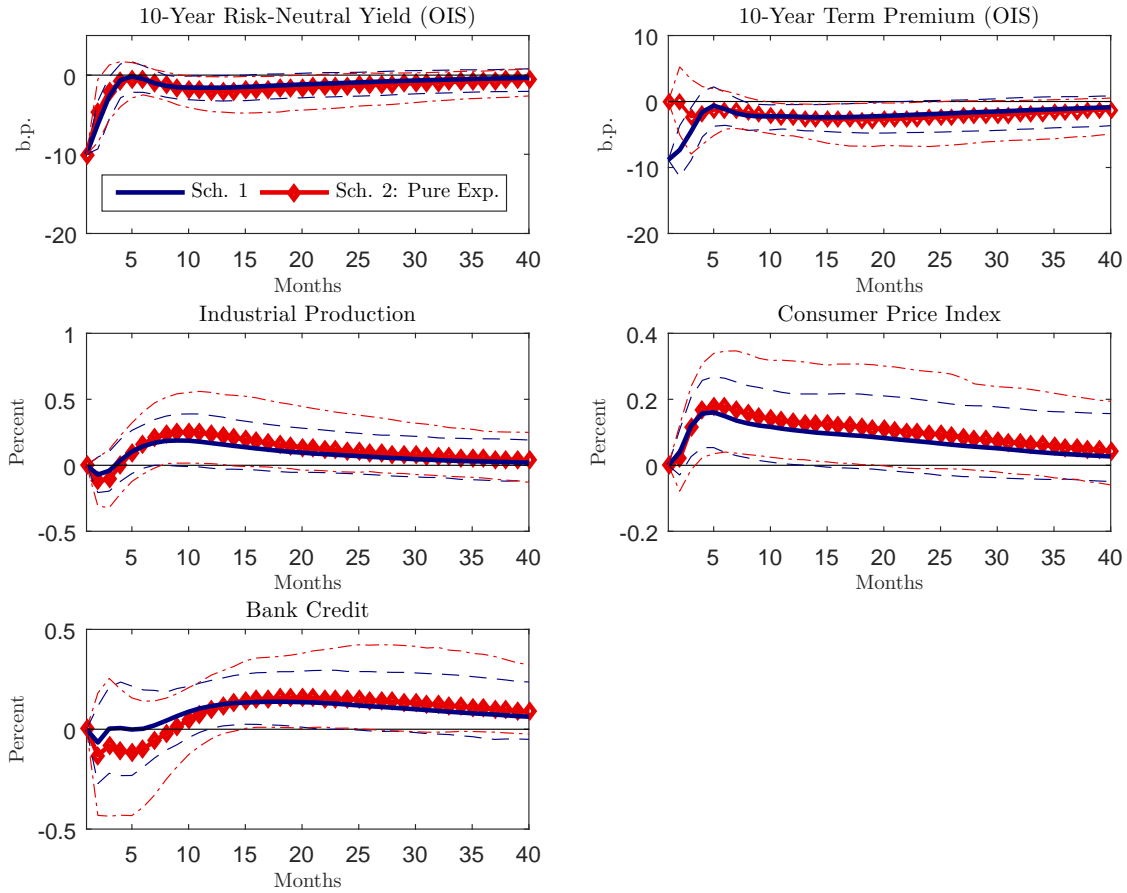
Here, I present the results from models 2 and 3, which control for bank credit and real exchange rates respectively, using the OIS-augmented decomposition of the 10-year yield. As for model 1, these results also strengthen when a longer sample period (November 2008 to December 2015) is used, and are also robust to the use of different maturity yields and different yield curve decompositions. These robustness exercises are presented in Lloyd (2017c, Appendix C.2).

Figures 4.5 and 4.6 present the impulse response functions for model 2, which controls for bank credit, estimated with the OIS-augmented decomposition of the 10-year Treasury yield. The results lend further support to those from model 1. The signalling shock has significantly positive expansionary effects on industrial production and CPI. With scheme 1, the response of industrial production is significantly positive 7 to 9 months after the shock, peaking at 0.19% after 9 months. Similarly, under scheme 2, the pure expectations shock, industrial production's response is significantly positive 8 to 13 months after the shock, peaking at 0.25% after 10 months. The impulse response of CPI is also significantly positive 3 to 13 months after the shock under scheme 1, and 4 to 19 months after the shock under scheme 2. The peak responses are 0.16% after 5 months with scheme 1, and 0.18% after 5 months with scheme 2.

Bank credit significantly increases with a lag in response to the expansionary signalling shock. Its impulse response is significantly positive 12 to 21 months after the shock with scheme 1, and 15 to 25 months after the shock under scheme 2. The peak bank credit response is 0.14% after 18 months, and 0.16% after 20 months under schemes 1 and 2 respectively.

Figure 4.6 demonstrates that the impulse responses of industrial production and CPI to the portfolio rebalancing shock remain insignificantly different from zero when the VAR is extended to include bank credit as a control variable. As well as being statistically insignificant, the peak responses of industrial production are smaller than their counterparts from figure 4.5. The peak industrial production responses to the portfolio rebalancing shock are 0.07% after 3 and 7 months under schemes 1 and 2 respectively, less than half of the peak responses to signalling shocks. Comparing these figures to the peak responses of industrial production to the signalling shock

Figure 4.5: Impulse Response Functions to a Signalling Shock for Model 2 using Sign Restriction Schemes 1 & 2 and the OIS-Augmented Decomposition of the 10-Year Treasury Yield



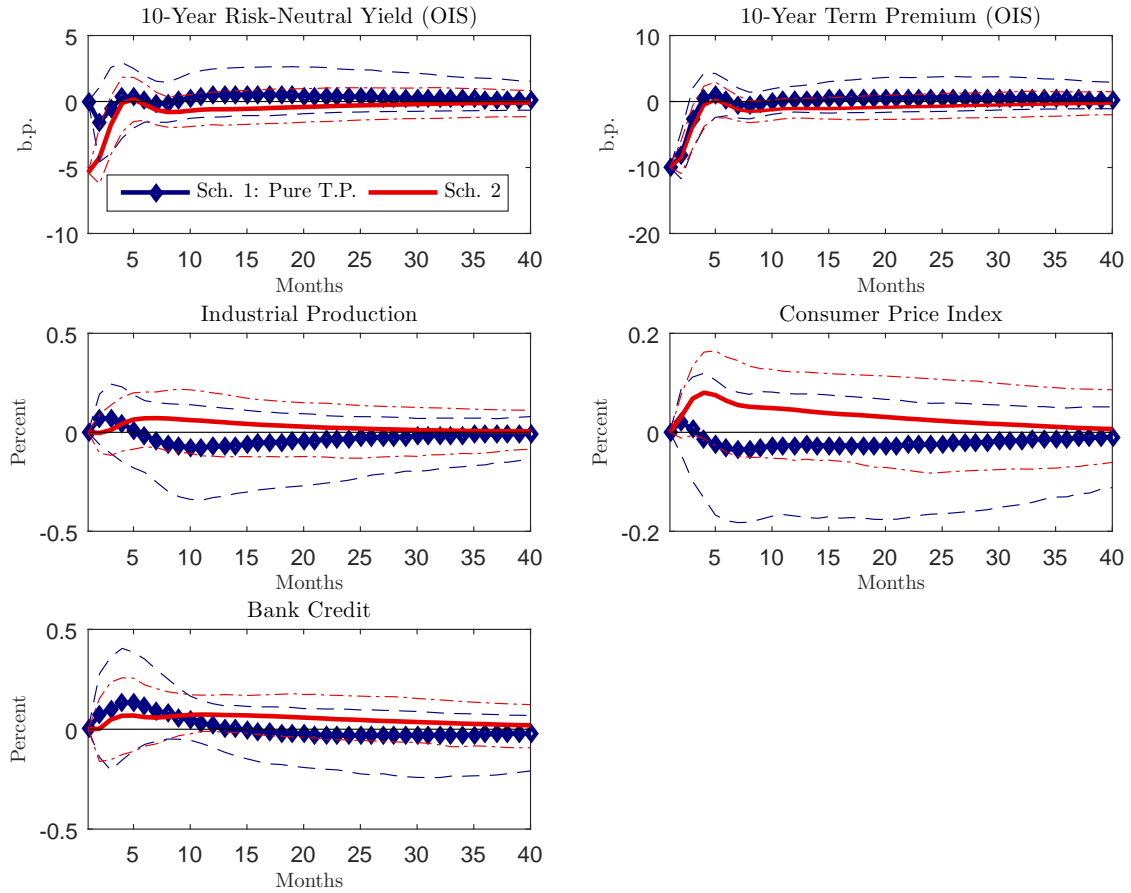
Note: The plots show the impulse response functions to a signalling shock, normalised to represent a 10 basis point fall in the 10-year risk-neutral yield, estimated using the OIS-augmented decomposition (Lloyd, 2017a). The VAR is estimated using data from November 2008 to April 2013. The two sign restriction schemes are detailed in table 4.8. The bold lines denote median impulse response draws. The thin dashed lines represent the 5% and 95% confidence intervals around median impulse responses, constructed using a residual-based block bootstrap.

indicates that the signalling shock explains around three-quarters of the total peak industrial production effect due to the long-term interest rate shocks.

The VAR results also indicate that portfolio rebalancing shocks had smaller expansionary effects on bank credit than signalling shocks. The peak bank credit increase is 0.13% after 5 months and 0.07% after 11 months under schemes 1 and 2 respectively. Under both schemes, the response of bank credit is insignificantly different from zero at all horizons.

The impulse response functions for model 3, which includes the real exchange rate, are presented in figures 4.7 and 4.8. Figure 4.7 demonstrates that the signalling shock has a significantly positive and lagged effect on industrial production and CPI under scheme 2. This pure expectations shock generates a significantly positive increase in industrial production 6 to 10 months after the shock, peaking at 0.47% after 7 months. Following the same shock, the peak CPI response is 0.25% after 4 months, and the responses are significantly positive from 3 to 6 and 13 to 39 months after the shock. As in models 1 and 2, the peak industrial production response following a signalling shock is smaller under scheme 1 than scheme 2. In model 3 with scheme 1, the peak is 0.20% after 8 months, although this is statistically insignificant using 5% and 95% confidence intervals. Nevertheless, the response of CPI to a signalling shock under scheme 1 is

Figure 4.6: Impulse Response Functions to a Portfolio Rebalancing Shock for Model 2 using Sign Restriction Schemes 1 & 2 and the OIS-Augmented Decomposition of the 10-Year Treasury Yield



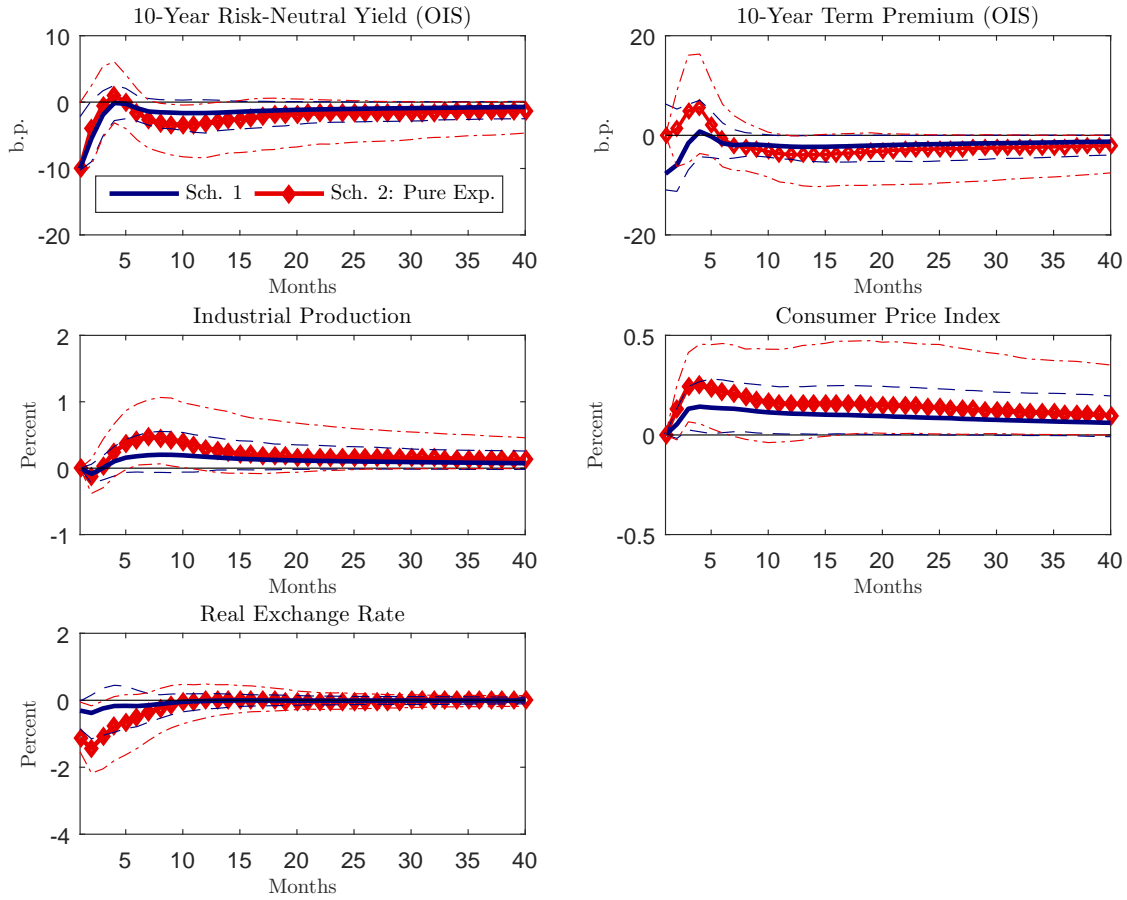
Note: The plots show the impulse response functions to a portfolio rebalancing shock, normalised to represent a 10 basis point fall in the 10-year term premium, estimated using the OIS-augmented decomposition (Lloyd, 2017a). The VAR is estimated using data from November 2008 to April 2013. The two sign restriction schemes are detailed in table 4.8. The bold lines denote median impulse response draws. The thin dashed lines represent the 5% and 95% confidence intervals around median impulse responses, constructed using a residual-based block bootstrap.

significantly positive 3 to 25 months after the shock, peaking at 0.14% after 4 months.

Under scheme 2, the peak response of industrial production to a signalling shock is the highest of all three models at 0.47%. This is likely to be explained by the response of the real exchange rate, which depreciates significantly on impact, and for the subsequent two months. This depreciation could stimulate industrial production through a trade balance channel. Interestingly, under scheme 1 which allows the term premium to respond to the signalling shock contemporaneously, the depreciation of the real exchange rate is not significant at any horizon, except on impact, and the response of industrial production is quantitatively smaller. This indicates that the responsiveness of the term premium to unconventional monetary policy announcements may have different implications for exchange rate movements than changes in expected future interest rates.

Figure 4.8 demonstrates that the impulse responses of industrial production and CPI to a portfolio rebalancing shock are not significantly positive at any horizon when the VAR is extended to include the real exchange rate as a control variable. In response to the portfolio rebalancing shock under both restriction schemes, the real exchange rate depreciates significantly

Figure 4.7: Impulse Response Functions to a Signalling Shock for Model 3 using Sign Restriction Schemes 1 & 2 and the OIS-Augmented Decomposition of the 10-Year Treasury Yield



Note: The plots show the impulse response functions to a signalling shock, normalised to represent a 10 basis point fall in the 10-year risk-neutral yield, estimated using the OIS-augmented decomposition (Lloyd, 2017a). The VAR is estimated using data from November 2008 to April 2013. The two sign restriction schemes are detailed in table 4.8. The bold lines denote median impulse response draws. The thin dashed lines represent the 5% and 95% confidence intervals around median impulse responses, constructed using a residual-based block bootstrap.

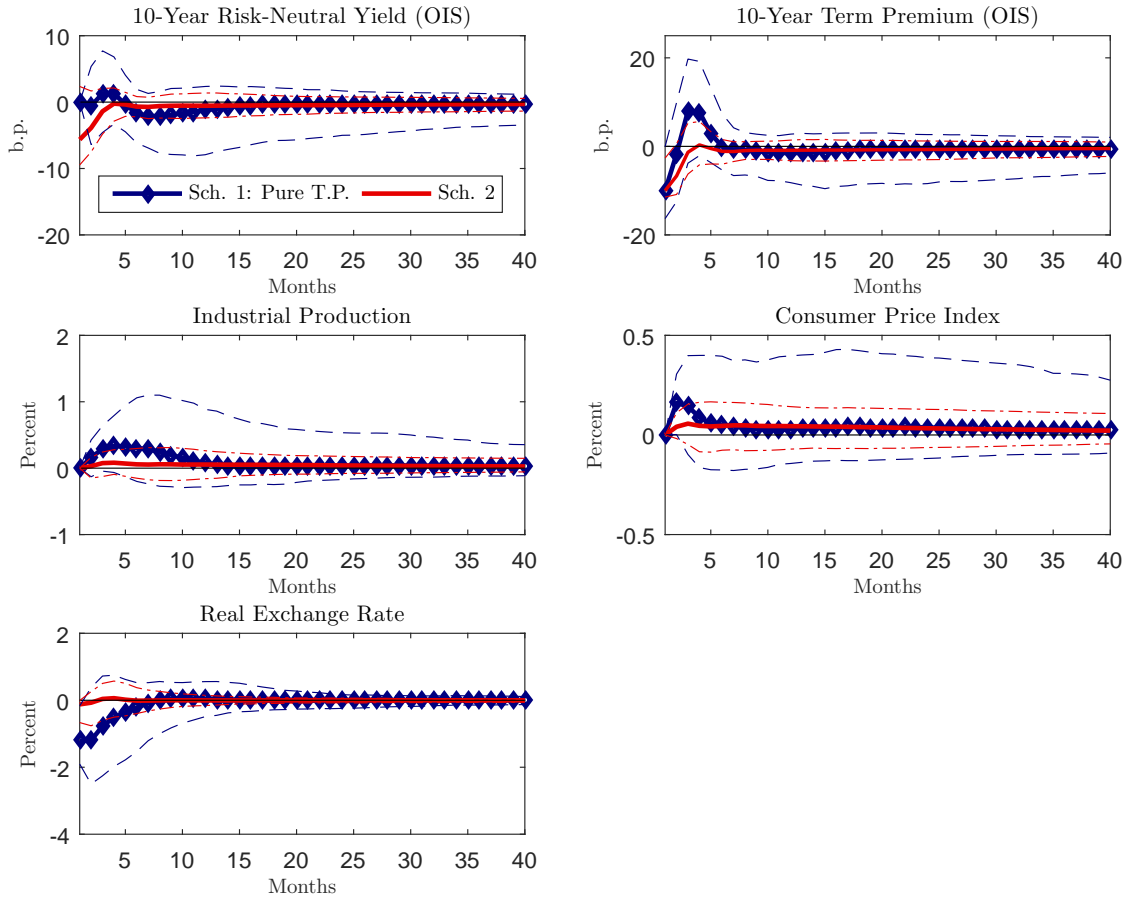
on impact as the identification restrictions impose, but its response thereafter is insignificantly different from zero.

The robustness exercises here, together with those in Lloyd (2017c, Appendix C), strongly indicate that shocks to expected future short-term interest rates have larger and more significant effects on industrial production and CPI than shocks to term premia. Thus, the signalling channel is associated with the majority of total industrial production and CPI effects attributed to shocks to longer-term interest rates over the November 2008 to April 2013 sample period. Across the three models the signalling channel is associated with around two-thirds to three-quarters of the industrial production effects attributed to the two channels.

4.6 Conclusion

In response to financial turmoil and the ELB for the short-term nominal policy rate, the Fed enacted a number of ‘unconventional’ monetary policies, notably LSAPs and forward guidance. Both can transmit to the real economy via longer-term interest rates through two sub-channels: signalling and portfolio rebalancing. In this chapter, I ask: through which of these channels

Figure 4.8: Impulse Response Functions to a Portfolio Rebalancing Shock for Model 3 using Sign Restriction Schemes 1 & 2 and the OIS-Augmented Decomposition of the 10-Year Treasury Yield



Note: The plots show the impulse response functions to a portfolio rebalancing shock, normalised to represent a 10 basis point fall in the 10-year term premium, estimated using the OIS-augmented decomposition (Lloyd, 2017a). The VAR is estimated using data from November 2008 to April 2013. The two sign restriction schemes are detailed in table 4.8. The bold lines denote median impulse response draws. The thin dashed lines represent the 5% and 95% confidence intervals around median impulse responses, constructed using a residual-based block bootstrap.

— signalling or portfolio rebalancing — did US unconventional monetary policy have the most expansionary effects on output? I conclude that signalling shocks exerted a more powerful influence on US industrial production and consumer prices than portfolio rebalancing effects.

The findings are important because the two channels offer differing implications for monetary policy. Signalling implies that policy is most effective when investors anticipate that the central bank will keep short-term interest rates lower for longer. Portfolio rebalancing requires the large scale of LSAPs in order to generate sufficient portfolio substitution to reduce yields and influence activity in the real economy during periods of sizeable financial turmoil.

In reaching my conclusion, I offer solutions to two challenges: the yield curve decomposition associated with signalling and portfolio rebalancing, and the identification of signalling and portfolio balance shocks.

In response to the first challenge, I use the OIS-augmented decomposition proposed in chapter 3. This is the first study to use this decomposition to assess the effects of macroeconomic policy at the ELB. I provide further evidence in support of the conclusions in chapter 3: the OIS-augmented GADTSM provides risk-neutral yields that closely align with the interest rate

expectations implied by FFFs and surveys. Using these daily frequency decompositions of the yield curve, I find that the signalling effects of unconventional monetary policy were particularly powerful at horizons in excess of 2 years. Monetary policy can have powerful effects on longer-term interest rate expectations. Signalling effects explain around two-thirds to three-quarters of the falls in 10-year bond yields on unconventional monetary policy announcement days. The result highlights the economic importance of the signalling effects of unconventional monetary policy announcements, implying that clear communication is likely to be an important determinant of the macroeconomic impact of central bank balance sheet normalisation.

In response to the second challenge, I identify shocks using two combinations of sign and zero-impact restrictions within an SVAR. I find that the signalling channel is associated with around two-thirds to three-quarters of the total peak effects of long-term interest rate shocks on US industrial production between November 2008 and April 2013.

The main conclusion of this chapter is that, as a result of reductions in longer-term interest rates between November 2008 and April 2013, the greatest benefits for industrial production and consumer prices were attributable to signalling. Thus, my findings suggest that current and future unconventional monetary policy action by the Fed may reap greater real economic benefits if combined with clear communication about the future short-term interest rate path.

Chapter 5

Long-Term Interest Rates, International Risk Sharing and Global Macroeconomic Spillovers

5.1 Introduction

With short-term nominal interest rates at their effective lower bound (ELB) following the 2007-2008 financial crisis, monetary policymakers sought to stimulate economic activity, *inter alia*, by reducing longer-term interest rates with ‘unconventional’ monetary policies. Although these policies have had domestic effects (chapter 4), they have also drawn attention to the global implications of long-term interest rate movements (Fratzscher, Lo Duca, and Straub, 2016). Understanding the role of long-term interest rates in the global transmission of shocks is the primary subject of this chapter. Specifically, I ask: to what extent, and through which mechanisms, does conventional US monetary policy spill over to other advanced economies through longer-term interest rates?

This is an important question in light of recent evidence suggesting that the Mundellian trilemma — which states that it is impossible for an open economy to simultaneously have a fixed exchange rate, free international capital movement and independent monetary policy — is, in fact, a dilemma — a choice between independent monetary policy and free capital mobility (Rey, 2014). Advanced economies with floating exchange rates are exposed to US monetary policy through its influence on the ‘global financial cycle’, which describes the strong correlation between financial market prices and international capital flows (Passari and Rey, 2015). Although recent work has shown US monetary policy to influence the global financial cycle (e.g. Miranda-Agrippino and Rey, 2015; Dedola, Rivolta, and Stracca, 2016; Rey, 2016), further research is required to understand the channels through which US monetary policy exerts international spillover effects. In this chapter, I assess the extent to which US monetary policy has global spillover effects through its influence on global bond markets and the term structure of interest rates in other advanced economies.

I, again, draw on the canonical decomposition of longer-term interest rates into expectations of future short-term interest rates and term premia, which is salient because the two sub-components have differing policy implications. If US monetary policy predominantly exerts spillover effects through changes in expected future short-term interest rates, monetary poli-

cymakers in other advanced economies may attenuate these effects by clearly communicating through policies such as forward guidance. In contrast, spillover effects that work through term premia motivate a careful focus on international capital flows and risk premia.

I study the international transmission of US monetary policy shocks from both a theoretical and an empirical standpoint. The two-country, two-good, micro-founded model I present has three essential elements.

First, and most importantly, the model allows for endogenous portfolio choice across a range of internationally traded assets: country-specific equity, short-term bonds, and long-term bonds. To incorporate these six assets endogenously, I use the solution method for optimal asset allocation in international macroeconomic models proposed by Devereux and Sutherland (2010, 2011). Unlike existing applications of this solution algorithm, I extend it to investigate multiple bond maturities simultaneously. Within the model, the responses of interest rates to macroeconomic shocks will reflect portfolio adjustment by international investors.

Second, the model assumes that consumers have a greater preference for locally-produced goods, generating consumption home bias. This is consistent with empirical evidence that the majority of consumption consists of locally-produced goods (Kollmann, 2006). Consumption home bias implies that the real exchange rate fluctuates in response to shocks, generating an incentive for investors to take positions that hedge real exchange rate variation.

Third, the model incorporates nominal rigidities, admitting the study of monetary policy. Nominal rigidities are important for generating realistic asset positions within the model. Engel and Matsumoto (2009) show that equity home bias can arise in a two-country model with nominal rigidities because price stickiness generates a negative covariance between relative Home non-financial income (i.e. labour income) and relative Home equity returns.¹ Because of this negative correlation, Home equity provides a good hedge against non-financial income risk.²

The model provides two sets of novel contributions. First, because the model includes short *and* long-term bonds, it provides insights about the different roles of the two asset classes in international portfolios. In the model's equilibrium, international asset positions are a function of all sources of risk in the economy. Nevertheless, I find that short-term bond holdings predominantly insure against the immediate response of real exchange rates to macroeconomic shocks, as well as the reaction of monetary policy. For instance, following a negative Home productivity shock, the Home real exchange rate appreciates, reducing the relative return on Foreign short-term bonds in Home consumption units. In this state of nature, Home marginal utility is relatively high, so Home short-term bonds will *ex ante* provide a good hedge against this risk. Long-term bond holdings are sensitive to macroeconomic shocks that change their relative resale value, related to expected inflation and real exchange rate movements in particular. Unlike short-term bonds, Foreign long-term bonds provide a better *ex ante* hedge for

¹In Engel and Matsumoto (2009), domestic firms who cannot cut prices following a positive Home productivity shock are forced to economise on labour, reducing relative Home non-financial income, while relative Home profits and stock returns increase.

²Heathcote and Perri (2013) provide an alternative explanation for equity home bias in a flexible price two-country model which includes productive capital. In their model, equity home bias can arise from the dynamics of investment. For instance, following a positive Home productivity shock, relative Home labour income increases with wages. Yet Home firms invest more, inducing a fall in the relative price of Home capital, reducing relative Home profits and relative Home equity returns. Thus, the models of both Engel and Matsumoto (2009) and Heathcote and Perri (2013) generate a negative covariance between relative earnings and relative stock returns, although the underlying mechanisms differ. In this study, the mechanism that generates equity home bias resembles Engel and Matsumoto (2009) because I do not include capital in the model.

negative Home productivity shocks. Following such a shock, heightened and persistent relative Home inflation will increase the relative value of Foreign long-term bonds. As in Coeurdacier and Gourinchas (2016), equity mainly insures against non-financial income risk after controlling for bond returns.

Second, the model provides insights about the international transmission of conventional US monetary policy shocks through the term structure of interest rates that align with empirical evidence from an event study into the effects of US conventional monetary policy surprises on bond yields in other advanced economies. Within the model, a surprise tightening of US monetary policy generates an immediate increase in bond yields in the US and other advanced economies as a consequence of global portfolio rebalancing, suggesting that movements in interest rates amplify the global spillover effects of US monetary policy.

To extend the event study and shed further light on the global transmission of US monetary policy, I use the decomposition of long-term interest rates into interest rate expectations and term premia from chapter 3. I find that the spillover effects of US monetary policy are closely associated with changes in expectations: a surprise tightening of US monetary policy immediately increases investors' expectations of future short-term interest rates in the US and other advanced economies. However, following the same monetary policy surprise, term premia fall, especially at longer horizons, suggesting that portfolio rebalancing by investors serves to attenuate global spillover effects.

The remainder of this chapter is structured as follows. After a brief review of the related literature, section 5.2 presents the stylised facts regarding the size and composition of bilateral international asset positions amongst advanced economies. Importantly, the model is able to account for many of these stylised facts in reasonable regions of the parameter space. The model and its solution algorithm are laid out in section 5.3. Section 5.4 presents the model calibration and discusses the role of the different assets in international risk sharing, emphasising the differences between short and long-term bonds. The global spillover effects of US monetary policy and its transmission through longer-term interest rates are discussed from both a theoretical and an empirical standpoint in section 5.5. Section 5.6 concludes.

Literature Review There has been a vast literature studying the growth in cross-border asset trade in the last three decades (e.g. Lane and Milesi-Ferretti, 2001, 2007; Coeurdacier and Rey, 2013; Gourinchas and Rey, 2014). Theoretical research has sought to understand the role different assets play in hedging different sources of risk. A long literature studying international risk sharing (e.g. Cole and Obstfeld, 1991; Backus, Kehoe, and Kydland, 1992, 1994; Stockman and Tesar, 1995; Corsetti, Dedola, and Leduc, 2008b) has provided insights about the determinants of the size of international asset positions. Recent computational advances by Devereux and Sutherland (2010, 2011) and Tille and van Wincoop (2010) have enabled a more detailed study of the *composition* of international portfolios, admitting endogenous portfolio choice amongst multiple assets. Building on this, papers have assessed the role played by country-specific equity *and* short-term bonds in international risk sharing (e.g. Devereux and Sutherland, 2008; Coeurdacier, Kollmann, and Martin, 2009, 2010; Coeurdacier and Gourinchas, 2016).³ This chapter builds on this literature by studying the role of country-specific short *and* long-term

³Additionally, Heathcote and Perri (2013) study an environment with country-specific equity, and Benigno and Küçük (2012) present a model with country-specific short-term bonds.

bonds, as well as country-specific equity, in order to understand the influence of international bond positions on the global financial cycle.

The primary contribution of this chapter is to study the mechanisms through which US monetary policy exerts global spillover effects via long-term interest rates. As such, this chapter is related to a literature studying the global financial cycle and the influence of US monetary policy on it (e.g. Rey, 2014, 2016). Passari and Rey (2015) present evidence on the existence of a global financial cycle in gross cross-border flows, asset prices and leverage, motivating research by Miranda-Agrippino and Rey (2015) documenting the effects of US monetary policy on the global financial cycle. Dedola et al. (2016) study the global macroeconomic effects of US monetary policy, emphasising the differences in transmission to advanced and emerging economies. This study builds on this literature by studying a particular aspect of the global financial cycle for advanced economies, related to international bond portfolios and the term structure of interest rates.

Because the empirical setup is able to distinguish between two sub-components of longer-term interest rates — interest rate expectations and term premia — it can shed light on the differential roles for central bank communication and international portfolio adjustment in global macroeconomic stabilisation. The empirical results presented here rely on the overnight indexed swap (OIS) augmented Gaussian Affine Dynamic Term Structure Model (GADTSM) from chapter 3, which enables the estimation of interest rate expectations and term premia at a daily frequency. The results in chapters 3 and 4 pertain to the US only. This is the first study to apply the OIS-augmented GADTSM to a broader set of countries, and is made possible because OIS rates offer a globally-comparable market-based measure of interest rate expectations in advanced economies (see chapter 2). Because this chapter considers the evolution of bond yields in multiple economies, it goes some way to uncovering information about the interaction between the term structure of interest rates and exchange rates. Stavrageva and Tang (2016) have recently empirically investigated the comovement of interest rates and exchange rates among advanced economies. They find that the relationship between short-term interest rates and exchange rate changes is primarily driven by expectations of future short-term interest rates, while the comovement of longer-term interest rates and exchange rates is more sensitive to term premia. My empirical results suggest that US monetary policy announcements influence these comovements, as I find that US monetary policy announcements have stronger effects on monetary policy expectations over shorter horizons, but larger effects on term premia at longer horizons.

5.2 Stylised Facts

The theoretical model presented in section 5.3 has a rich financial market structure, accounting for country-specific equity, and short and long-term bonds. In this section, I present empirical evidence about the size and composition of international portfolios, motivating the need to consider this set of assets within a theoretical framework, as all play a role in international asset portfolios, and providing targets that guide model calibration.

5.2.1 Size of International Asset Portfolios

Throughout the chapter, I focus on a sub-sample of advanced economies with floating exchange rates: Australia, Canada, France, Germany, the UK, and the US. This choice is motivated by

a number of factors, primary amongst which is the availability of daily frequency zero-coupon government bond yield data over a sufficiently long sample. This data is necessary in section 5.5 where I empirically decompose longer-term interest rates into expectations of future interest rates and term premia using the method from chapter 3. Notwithstanding this, there are economic reasons to consider this sub-sample of countries. First, these countries comprise a large fraction of international asset trade; the combined total international asset position of the six economies averaged 42% of total world international portfolio holdings over the 2001-2014 period.⁴ Second, these economies individually comprise a large fraction of world international asset positions, as shown in figure 5.1. Here, I plot the average size of countries' international asset position as a percentage of total world portfolio investment for the 2001-2014 period for the twenty economies with the largest foreign asset portfolios, ranked in order of size. All six economies included in this study are in the world top twenty, and four of them are in the world top six.⁵ Third, these six economies are not 'offshore financial centres'.⁶ This is important because the international asset positions attributed to offshore financial centres are not likely to reflect the ultimate country of asset ownership.⁷

I study bilateral financial flows between these economies, with the US acting as the base country in all cases.⁸ This is motivated by the US's central role in the international financial system. The US dollar is an important funding currency for international banking and is widely used by investment fund managers. Figure 5.1 illustrates the US's important role in the international financial system, showing that it accounted for almost 18% of total world foreign portfolio investment between 2001 and 2014, around 9 percentage points more than the second highest economy, the UK.

Moreover, the US is an important counterparty to cross-border asset holdings for the remaining five countries considered. US investors hold significant quantities of assets issued in Australia, Canada, France, Germany and the UK, as shown in figure 5.2. Here I present average total US foreign portfolio investment holdings by country of asset issuer as a percentage of US total foreign portfolio investment for the 2001-2014 period, demonstrating how exposed US asset holders are to assets issued in other world economies. The figure plots the top 20 countries ranked by size of holdings, illustrating that US holdings of foreign assets issued in the UK, Canada, France, Germany and Australia are amongst the top nine countries overall.

Similarly, a large fraction of foreign assets held by Australian, Canadian, French, German and UK investors are US-issued. Figure 5.3 plots average foreign portfolio investment holdings of Australia, Canada, France, Germany and the UK by country of asset issuer as a percentage of total foreign portfolio investment for 2001-2014, indicating how exposed investors in these countries are to assets issued elsewhere. For three out of the five countries — Australia, Canada and the UK — holdings of US-issued assets comprise the largest fraction of the total foreign

⁴This statistic is calculated using annual data from the Coordinated Portfolio Investment Survey (CPIS) by the International Monetary Fund (IMF). See appendix A.1 for more details on this data source.

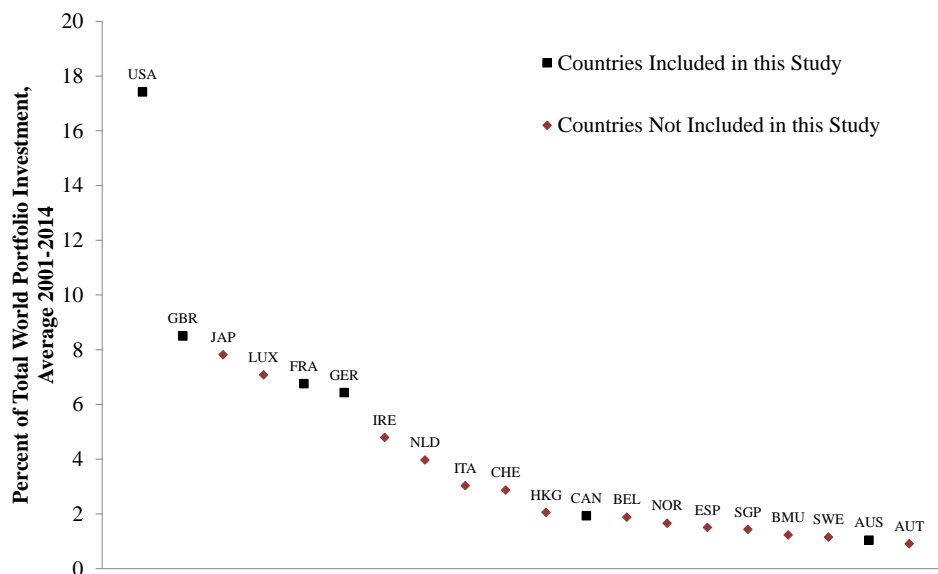
⁵Although Japan individually accounts for the third largest fraction of world asset holdings, I omit this from my study due to its unique economic conditions over the last two decades following the collapse of the Japanese asset price bubble in the 1990s.

⁶I use the IMF classification of offshore financial centres, provided in the following report: www.imf.org/external/np/pp/eng/2008/050808.pdf.

⁷Amongst the top twenty economies in figure 5.1, six are classed as offshore financial centres by the IMF: Bermuda, Hong Kong, Ireland, Luxembourg, Singapore and Switzerland. The remaining countries in figure 5.1 are omitted from this study because of a lack of detailed daily frequency zero-coupon government bond yield data.

⁸In the two-country theoretical model presented in section 5.3, the US is the 'Foreign' country.

Figure 5.1: Total Foreign Portfolio Investment Holdings as a Percentage of Total World Portfolio Investment, Top 20 World Economies, Average 2001-2014



Notes: The average value of total foreign portfolio investment (holdings) for a country as a percentage of total world portfolio investment over the 2001-2014 period (annual data) for the 20 world economies with the largest foreign asset portfolios, ranked in order of size. Economies denoted with black squares are those studied in this chapter — United States (USA), United Kingdom (GBR), France (FRA), Germany (GER), Canada (CAN), Australia (AUS). Economies denoted with diamonds are not included in this study, and are comprised of offshore financial centres (as defined by the IMF) — Luxembourg (LUX), Ireland (IRE), Switzerland (CHE), Hong Kong (HKG), Singapore (SGP), Bermuda (BMU) — economies lacking publicly available daily zero-coupon government bond yield data — the Netherlands (NLD), Italy (ITA), Belgium (BEL), Norway (NOR), Spain (ESP), Sweden (SWE), Austria (AUT) — and Japan (JAP), which is omitted from this study due to its unique economic conditions over the last two decades. *Data Source:* Coordinated Portfolio Investment Survey by the IMF and author's own calculations.

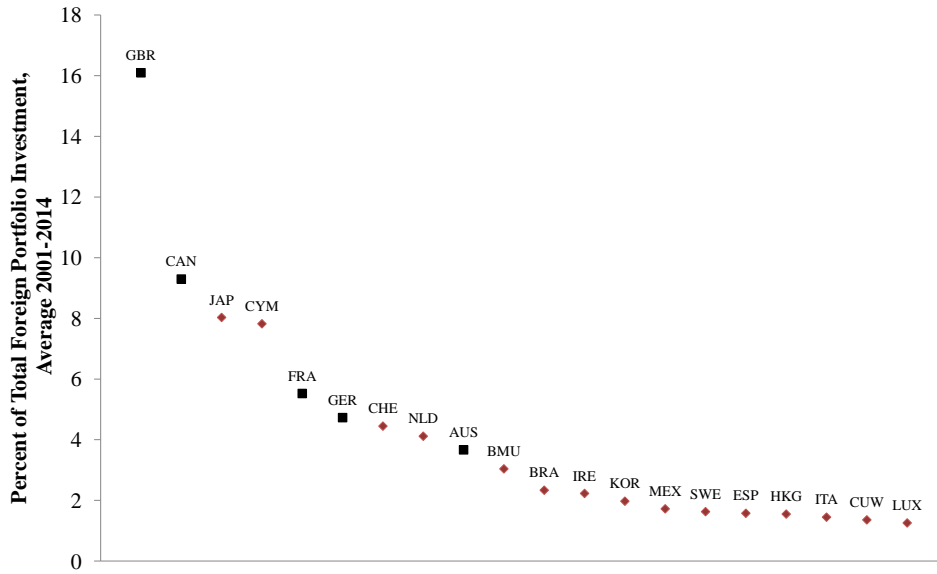
asset portfolio. Moreover, their US-issued asset holdings are at least four times larger than holdings of foreign assets from the second-highest country. For these economies, the US is by far the primary counterparty in international asset trade. French and German portfolios are, unsurprisingly, more Eurozone-oriented. US-issued assets comprise the fourth and third largest shares of total international asset positions in these two countries, respectively. Nevertheless, although their foreign asset portfolios are not as concentrated towards a single country, US-issued assets remain a sizeable fraction of their international portfolio (at least 9%).

Although the US is an important counterparty for Australia, Canada, France, Germany and the UK, figure 5.3 shows that a sizable quantity of bilateral portfolio investment occurs between these five countries too, further motivating the selection of these countries.

5.2.2 Composition of Bilateral Asset Portfolios

Although the size of cross-border financial flows is an important feature of the international financial system, the composition of cross-border flows is the primary focus of this chapter. To date, empirical studies of international portfolio composition have focused on foreign currency exposure (Lane and Shambaugh, 2010a,b) and the sectoral breakdown of international asset

Figure 5.2: US Foreign Portfolio Investment Holdings by Country of Asset Issuer as a Percentage of Total US Foreign Portfolio Investment, Top 20 World Economies, Average 2001-2014



Notes: The average value of US foreign portfolio investment (holdings) by country of asset issuer as a percentage of total US foreign portfolio investment over the 2001-2014 period (annual data) for the top 20 economies ranked in order of size. Economies denoted with squares are those studied in this chapter — United States (USA), United Kingdom (GBR), France (FRA), Germany (GER), Canada (CAN), Australia (AUS). Economies denoted with diamonds are excluded from this study — Japan (JAP), Cayman Islands (CYM), Switzerland (CHE), the Netherlands (NLD), Bermuda (BMU), Brazil (BRA), Ireland (IRE), Republic of Korea (KOR), Mexico (MEX), Sweden (SWE), Spain (ESP), Hong Kong (HKG), Italy (ITA), Curaçao (CUW), Luxembourg (LUX). *Data Source:* Coordinated Portfolio Investment Survey by the IMF and author's own calculations.

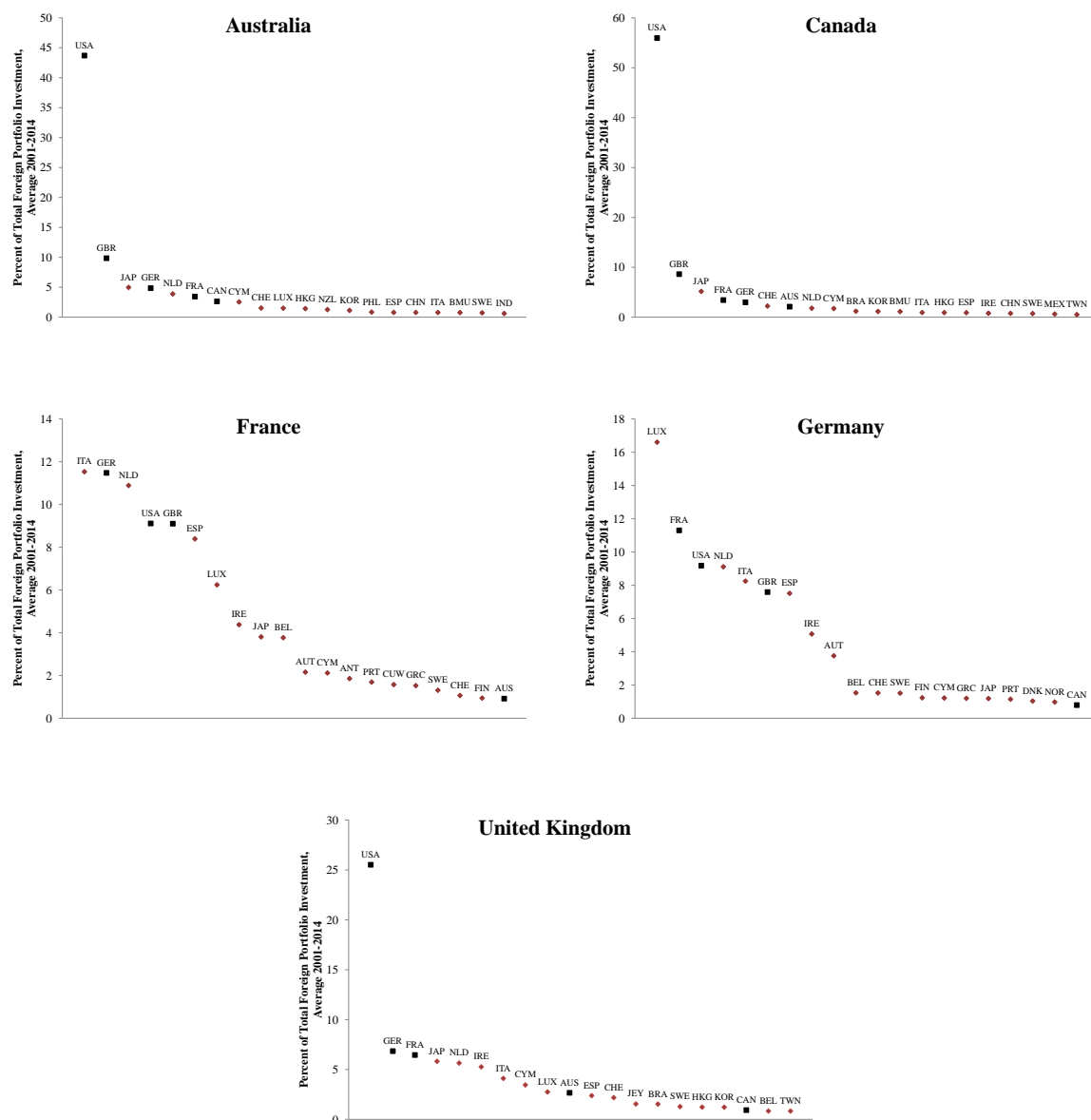
holdings (Galstyan, Lane, Mehigan, and Mercado, 2016). In this chapter, I study the composition of bilateral international asset positions by asset class.

Data I use data from the US Treasury International Capital (TIC) system, the US government's source on capital flows into and out of the United States.⁹ I use the annual holdings data from the TIC system reported by issuers and holders of US and foreign securities. This data is collated from annual surveys at the end of June each year, beginning in 2002, which collect holdings information by individual securities for private investors and are thus considered to be highly accurate.¹⁰ The dataset includes information on the stock of bilateral cross-border asset holdings between US residents (including US-based branches of firms headquartered in other countries) and foreign residents (including offshore branches of US firms). The dataset permits analysis along many dimensions, including currency, type of foreign holder, industry of issuer, and, of particular interest in this study, country of foreign holder, security type and maturity. I exclude holdings of US-issued assets by foreign official institutions (FOIs), because the incentives for private investors for holding US-issued assets may differ from those of FOIs.

⁹See appendix A.1 for more details on this data source.

¹⁰Unlike the survey collection underlying the CPIS by the IMF, reporting for the TIC survey is compulsory, and significant penalties can be imposed for a failure to report.

Figure 5.3: Foreign Portfolio Investment Holdings by Country of Asset Issuer as a Percentage of Total Foreign Portfolio Investment for Australia, Canada, France, Germany and the United Kingdom, Top 20 World Economies, Average 2001-2014



Notes: The average value of Australian, Canadian, French, German and UK foreign portfolio investment (holdings) by country of asset issuer as a percentage of total foreign portfolio investment in each country over the 2001-2014 period (annual data). Each graph plots the top 20 economies ranked in order of size. Economies denoted with squares are those studied in this chapter — United States (USA), United Kingdom (GBR), France (FRA), Germany (GER), Canada (CAN), Australia (AUS). Economies denoted with diamonds are excluded from this study — Antilles (ANT), Austria (AUT), Belgium (BEL), Bermuda (BMU), Brazil (BRA), Cayman Islands (CYM), China (CHN), Curacao (CUW), Denmark (DNK), Finland (FIN), Greece (GRC), Hong Kong (HKG), India (IND), Ireland (IRE), Italy (ITA), Japan (JAP), Jersey (JEY), Korea (KOR), Luxembourg (LUX), Mexico (MEX), Netherlands (NLD), New Zealand (NZL), Norway (NOR), Philippines (PHL), Portugal (PRT), Spain (ESP), Sweden (SWE), Switzerland (CHE), Taiwan (TWN). *Data Source:* Coordinated Portfolio Investment Survey by the IMF and author's own calculations.

I use data on foreign holdings of US-issued equity, short-term debt and long-term debt. Equity is defined to include: common and preferred stock; all types of investment company shares, such as open and closed-end funds, money market mutual funds, and hedge funds; and interest in limited partnerships and other equity interests that may not involve stocks or shares.¹¹ Debt is defined to include US Treasury debt, US agency debt,¹² and corporate debt.¹³

Within the dataset, long-term debt includes all debt with *original term-to-maturity* in excess of one year, while short-term debt includes all debt with *original term-to-maturity* of one year or less. I use information on the remaining years to maturity of foreign private holdings of US long-term debt reported in the TIC data to redefine the classifications for short and long-term debt. I define long-term debt to include all debt with *remaining time-to-maturity* in excess of one year, while short-term debt includes all debt with *remaining time-to-maturity* of one year or less, to match model definitions in section 5.3.

Empirical Evidence The US TIC data highlights novel information about the composition of bilateral holdings of US-issued assets. Figure 5.4 presents the annual time series of holdings of US equity and debt as a percentage of the holding country’s nominal GDP.

Figure 5.4 emphasises three important empirical regularities. First, all five countries hold positive quantities of all classes of US-issued assets; they do not have net short positions in US-issued equity or debt. Second, with the exception of Canada, countries’ exposure to US-issued debt is roughly comparable in size to their exposure to US-issued equity. For example, in 2003, UK holdings of US-issued equity and US-issued debt were both 10% of UK GDP. This motivates the need to jointly consider the role of equity *and* bonds in international asset portfolios. Third, the majority of debt holdings are associated with long-term debt, supporting the separate concern for short and long-term bonds in this chapter. Countries’ holdings of US-issued short-term debt range from 0.52% to 3.36% of GDP, while holdings of US-issued long-term debt lie between 1.53% and 19.19% of GDP.

Figure 5.4 also illustrates that there are various dimensions of heterogeneity in the size and composition of US-issued asset holdings across countries. Most strikingly, there is heterogeneity in the size of a foreign countries’ exposure to different classes of US-issued assets. During the 2002-2015 period, Canada held the highest stock of US-issued equity of all five countries, with the stock as a percentage of its GDP over four times larger than the corresponding German and French figures in all years. Likewise, UK holdings of US-issued debt as a percentage of its GDP were over double the corresponding figure for Australia throughout the period.

Figure 5.4 further depicts time variation in bilateral asset holdings. Canadian, German and UK holdings of US-issued equity have increased in every year from 2009, around the time when monetary policy reached its ELB in many advanced economies.¹⁴ The stock of Canadian US-issued equity holdings in 2015 was 49%, almost three times its value in its 2009 trough, 18%. In contrast, UK holdings of US-issued debt, which increased from 10% of UK GDP in 2002 to 22% in 2009, have remained relatively stable since 2009; in 2015, UK holdings of US-issued debt

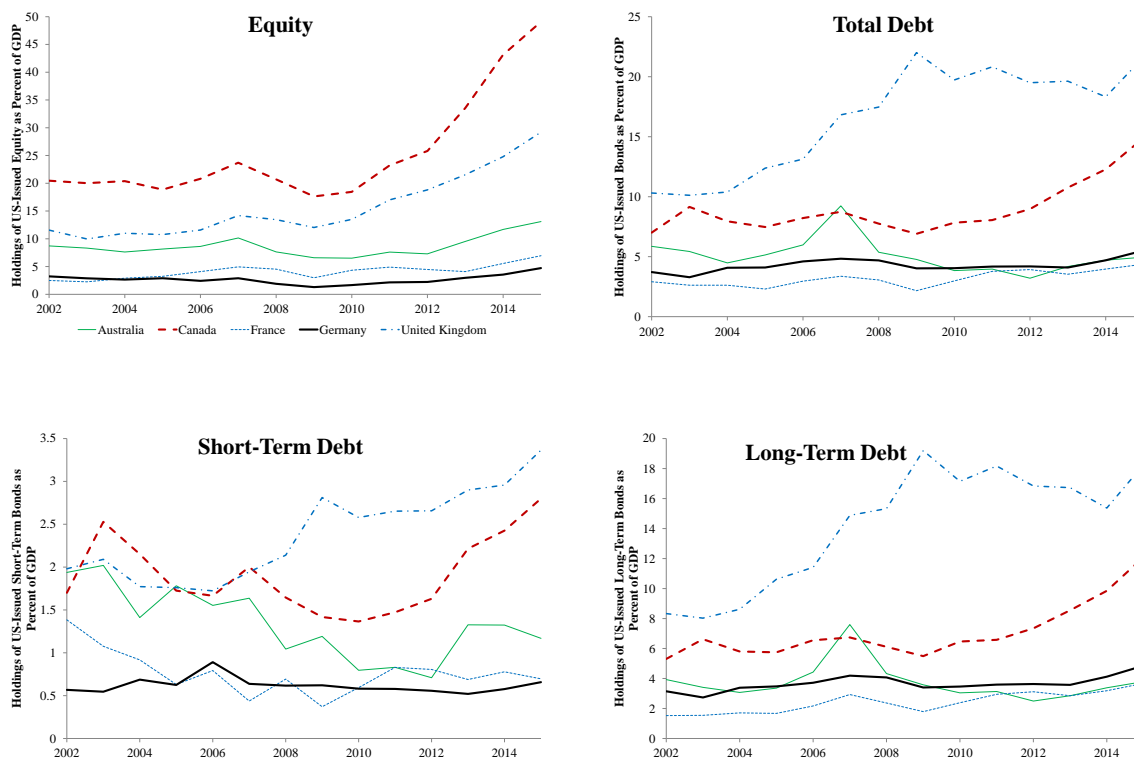
¹¹Because the data is collected on a security-by-security basis, double-counting is eliminated in the TIC system.

¹²Where agencies include US government agencies and corporations, as well as federally sponsored enterprises, such as the Federal National Mortgage Association.

¹³Corporate debt includes all non-Treasury and non-agency debt, such as corporate bonds, certificates of deposit and US municipal debt securities.

¹⁴The growth in the stock of US-issued equity held in Canada, Germany and the UK has exceeded the growth in US equity prices from their trough in February 2009 to 2015.

Figure 5.4: Total Private Holdings of US-Issued Assets in Australia, Canada, France, Germany and the UK as a Percentage of Nominal GDP, Annual 2002-2015



Notes: Total private holdings of US-issued assets (equity, debt, short-term debt, and long-term debt respectively) by Australia, Canada, France, Germany and the UK as a percentage of nominal GDP. The data presents stocks of asset holdings as of June 30 in each year, from 2002 to 2015. Equity includes common and preferred stock; all types of investment company shares; interest in limited partnerships and other equity interests that may not involve stocks or shares. Debt includes US Treasury debt, US agency debt, and corporate debt. Long-term debt includes all debt with remaining term-to-maturity in excess of one year, while short-term debt includes all debt with remaining term-to-maturity of one year or less. *Data Sources:* US Treasury International Capital (TIC) system, IMF and author's own calculations.

had fallen by just 0.4 percentage points from a peak in 2009. The lower right panel of figure 5.4 indicates that this variation in UK holdings of US-issued debt emanates from changes in the stock of long-term debt held; short-term debt holdings for all countries have varied within a small band ranging from 0.52% to 3.36% of GDP.

Table 5.1 presents summary statistics for the time series data in figure 5.4. The table displays the mean and standard deviation of the stock of US-issued asset holdings by asset class over three time periods: the whole 2002-2015 sample; a pre-ELB 2002-2008 sample; and a post-ELB 2009-2015 sample. The final column denotes the range of the average values across countries, used to guide the model calibration in section 5.4. Average holdings of US-issued equity range from 2.68% of GDP to 25.42% over the whole 2002-2015 sample period. Over the two sub-periods, only the upper end of this range is significantly affected; the average holdings of US-issued equity ranges from 2.70% of GDP to 20.71% during the 2002-2008 period and from 2.66% to 30.13% during the 2009-2015 period. A similar pattern emerges for the average value of total debt holdings, which range from 3.18% to 16.59% for the 2002-2015 period; in the 2002-2008 and 2009-2015 sub-samples the corresponding ranges are 2.84-12.95% and 3.53-20.23%, respectively.

Table 5.1: Mean (Standard Deviation) of Foreign Countries' Holdings of US-Issued Assets as a Percentage of Each Country's Nominal GDP by Asset Class, 2002-2015

Country	Australia	Canada	France	Germany	United Kingdom	Range
<u>Equity</u>						
2002-2015	8.68 (1.89)	25.42 (9.70)	4.13 (1.28)	2.68 (0.86)	15.67 (5.88)	2.68-25.42
2002-2008	8.46 (0.86)	20.71 (1.47)	3.49 (1.04)	2.70 (0.43)	<i>11.79 (1.50)</i>	2.70-20.71
2009-2015	8.90 (2.62)	30.13 (12.24)	4.76 (1.24)	2.66 (1.19)	<i>19.56 (6.12)</i>	2.66-30.13
<u>Total Debt</u>						
2002-2015	5.08 (1.43)	9.00 (2.23)	3.18 (0.66)	4.29 (0.54)	16.59 (4.41)	3.18-16.59
2002-2008	<i>5.93 (1.54)</i>	8.04 (0.73)	<i>2.84 (0.35)</i>	4.19 (0.56)	<i>12.95 (3.09)</i>	2.84-12.95
2009-2015	<i>4.23 (0.62)</i>	9.96 (2.85)	<i>3.53 (0.74)</i>	4.39 (0.54)	<i>20.23 (1.31)</i>	3.53-20.23
<u>Short-Term Debt</u>						
2002-2015	1.34 (0.42)	1.91 (0.45)	0.76 (0.25)	0.62 (0.09)	2.38 (0.53)	0.62-2.38
2002-2008	<i>1.63 (0.33)</i>	1.92 (0.33)	0.85 (0.31)	0.65 (0.11)	<i>1.92 (0.17)</i>	0.65-1.92
2009-2015	<i>1.05 (0.26)</i>	1.90 (0.57)	0.68 (0.16)	0.58 (0.04)	<i>2.85 (0.27)</i>	0.58-2.85
<u>Long-Term Debt</u>						
2002-2015	3.74 (1.23)	7.09 (1.89)	2.42 (0.71)	3.67 (0.51)	14.21 (3.97)	2.42-14.21
2002-2008	4.30 (1.54)	<i>6.12 (0.53)</i>	<i>1.99 (0.52)</i>	3.53 (0.51)	<i>11.03 (3.05)</i>	1.99-11.03
2009-2015	3.18 (0.44)	<i>8.05 (2.29)</i>	<i>2.85 (0.61)</i>	3.81 (0.52)	<i>17.39 (1.26)</i>	2.85-17.39

Notes: Mean (standard deviation) of the stock of US-issued equity and debt held in Australia, Canada, France, Germany and the UK (excluding FOI holdings) using annual data over three time periods: the whole 2002-2015 sample; a pre-effective lower bound (ELB) 2002-2008 sub-sample; and a post-ELB 2009-2015 sub-sample. These represent summary statistics for the time series data plotted in figure 5.4. Italicised entries denote sub-samples that are statistically different from one another using a difference-in-mean hypothesis test and a 5% significance level. *Data Source:* US Treasury International Capital (TIC) system, IMF and author's own calculations.

The majority of the change in total debt holdings is associated with long-term debt holdings, which range from 2.42% to 14.21% over the same period, and are 1.99-11.03% and 2.85-17.39% in the 2002-2008 and 2009-2015 sub-samples, respectively. Short-term debt holdings are relatively small — 0.62-2.38% of GDP for the whole sample.

To investigate whether differences between the pre and post-ELB samples are statistically significant, I carry out difference-in-mean significance tests for each country and each series. The significantly different sub-samples are italicised in table 5.1. For all asset classes, UK holdings of US-issued assets are significantly higher in the second sub-sample. In two cases — for total bonds and long-term bonds — the UK figure defines the maximum in the range. Similarly, French holdings of US-issued long-term debt, which define the minimum in the range, are significantly higher in the second sub-sample. Because there are significant differences across sub-samples for certain asset classes and countries that influence the ranges, I use the 2002-2008 summary statistics as the calibration targets for the theoretical model. For all asset classes, the ranges for the 2002-2008 sub-sample are the narrowest of all three sub-samples, so provide the most stringent targets for the model to match.

5.3 A Model of International Bond Positions

In this section, I present a micro-founded, two-country model of international asset portfolios, which I use to study the role of short and long-term debt in international risk sharing and the global transmission of US monetary policy through longer-term interest rates.

In the model, there are two countries — Home H and Foreign F (the US, denoted with an $*$). Each country is populated by a continuum of infinitely-lived, identical consumers with unit mass. Every period t , each individual in each country consumes a basket of Home and Foreign goods and supplies labour to domestic firms. Firms produce differentiated brands and are monopolistically competitive, facing nominal rigidities à la Calvo (1983). In each country, monetary policymakers set the short-term nominal interest rate according to a Taylor-type rule.

Households have access to six assets: country-specific short-term bonds, long-term bonds, and equity. That is, there are three dimensions to household portfolio choice: (i) country of asset issuance (Home or Foreign); (ii) type of asset (equity or bonds); (iii) maturity of asset (short or long-term bonds). Because of the limited asset availability in comparison to the numerous sources of uncertainty, international financial markets are incomplete.

5.3.1 Households

A representative Home household maximises discounted expected lifetime utility:

$$U_t = \mathbb{E}_t \sum_{s=0}^{\infty} \delta_{t+s} e^{\zeta_{C,t+s}} \left(\frac{(C_{t+s} - \gamma C_{t+s-1})^{1-\sigma} - 1}{1-\sigma} - e^{\zeta_{L,t+s}} \frac{L_{t+s}^{1+\frac{1}{\nu}}}{1+\frac{1}{\nu}} \right) \quad (5.1)$$

where $u(C_t) = \frac{(C_t - \gamma C_{t-1})^{1-\sigma} - 1}{1-\sigma}$ is the instantaneous consumption utility function, $\delta_t \in (0, 1)$ is the discount factor, C_t is Home consumption, γC_{t-1} is the stock of (external) habits in period t , $\gamma \in [0, 1]$ is the persistence of habit formation, $\sigma > 0$ is the coefficient of relative risk aversion, L_t is the labour supply of the Home household, and $\nu > 0$ is the Frisch elasticity of labour supply. $\zeta_{C,t}$ and $\zeta_{L,t}$ are household preference and labour supply shocks respectively. The Home shocks follow independent autoregressive processes of order one:

$$\zeta_{C,t} = \rho_C \zeta_{C,t-1} + \varepsilon_{C,t}, \quad \varepsilon_{C,t} \sim \mathcal{N}(0, \sigma_C^2), \quad \rho_C \in (0, 1), \quad \sigma_C > 0 \quad (5.2)$$

$$\zeta_{L,t} = \rho_L \zeta_{L,t-1} + \varepsilon_{L,t}, \quad \varepsilon_{L,t} \sim \mathcal{N}(0, \sigma_L^2), \quad \rho_L \in (0, 1), \quad \sigma_L > 0 \quad (5.3)$$

and similarly in the Foreign country with $\rho_C = \rho_C^*$, $\rho_L = \rho_L^*$, $\sigma_C = \sigma_C^*$ and $\sigma_L = \sigma_L^*$.

The discount factor δ_{t+s} is given by:

$$\delta_{t+s+1} = \delta_{t+s} \beta(C_{t+s}), \quad \delta_t = 1$$

It is well known that in open economy models with incomplete markets, such as this, equilibrium dynamics exhibit non-stationarity. Schmitt-Grohé and Uribe (2003) outline solutions to this problem for a small-open economy setup, including: endogenising the discount factor, portfolio adjustment costs, or a debt-elastic interest rate. In this chapter, I use an endogenous discount factor to induce stationarity, by assuming:¹⁵

$$\beta(C_t) = \omega C_t^{-\eta}$$

with $\eta \in [0, \sigma)$ and $\omega \bar{C}^{-\eta} \in (0, 1)$, where \bar{C} is the steady-state value of consumption, such that

¹⁵This solution is most readily applied to the model solution algorithm of Devereux and Sutherland (2010, 2011). Moreover, by endogenising the discount factor, I avoid making direct assumptions about the composition of portfolios that I set out to study endogenously in this model.

$\beta(C_t) \in (0, 1)$ and $\beta'(C_t) \leq 0$ for all t . The impact of δ_t on consumption is not internalised by individual decision makers, so that the discount factor depends on the average consumption in an economy, rather than an individual's own consumption.

C_t is the aggregate consumption basket of Home households, comprised of baskets of Home $C_{H,t}$ and Foreign $C_{F,t}$ goods, given by a constant elasticity of substitution (CES) index:

$$C_t \equiv \left[(a_H)^{\frac{1}{\phi}} (C_{H,t})^{\frac{\phi-1}{\phi}} + (1 - a_H)^{\frac{1}{\phi}} (C_{F,t})^{\frac{\phi-1}{\phi}} \right]^{\frac{\phi}{\phi-1}} \quad (5.4)$$

where $a_H \in [0, 1]$ represents the weight of Home goods in the aggregate consumption basket,¹⁶ and $\phi > 0$ denotes the elasticity of substitution between Home and Foreign goods — the ‘trade elasticity’. An expenditure minimisation problem for the Home household yields the aggregate Home consumer price index (CPI) corresponding to the basket in (5.4):

$$P_t = \left[a_H (P_{H,t})^{1-\phi} + (1 - a_H) (P_{F,t})^{1-\phi} \right]^{\frac{1}{1-\phi}} \quad (5.5)$$

where $P_{H,t}$ and $P_{F,t}$ are the Home country price indices for the baskets of Home and Foreign goods, respectively.

The baskets of Home $C_{H,t}$ and Foreign $C_{F,t}$ goods consumed by the Home households themselves comprise a continuum of differentiated brands, each with unit mass, that are imperfectly substitutable and are aggregated by the following CES indices:

$$C_{H,t} \equiv \left[\int_0^1 C_t(h)^{\frac{\theta-1}{\theta}} dh \right]^{\frac{\theta}{\theta-1}}, \quad C_{F,t} \equiv \left[\int_0^1 C_t(f)^{\frac{\theta-1}{\theta}} df \right]^{\frac{\theta}{\theta-1}} \quad (5.6)$$

where $C_t(h)$ denotes the Home agents' consumption of Home brand h , $C_t(f)$ denotes the Home agents' consumption of foreign brand f , and $\theta > 1$ denotes the elasticity of substitution between varieties — the ‘brand elasticity’. The associated price indices, again the result of an expenditure minimisation problem for the Home household, are:

$$P_{H,t} = \left[\int_0^1 P_t(h)^{1-\theta} dh \right]^{\frac{1}{1-\theta}}, \quad P_{F,t} = \left[\int_0^1 P_t(f)^{1-\theta} df \right]^{\frac{1}{1-\theta}} \quad (5.7)$$

where $P_t(h)$ is the Home country price of Home brand h and $P_t(f)$ is the Home country price of Foreign brand f .

Symmetric expressions for the Foreign household are presented in appendix D.1.1. Unless otherwise stated, the parameterisation across countries is symmetric (e.g. $\sigma = \sigma^*$, $\nu = \nu^*$, $\phi = \phi^*$, $\theta = \theta^*$, $\omega = \omega^*$, $\eta = \eta^*$ and $a_H^* = 1 - a_H$, where a_H^* denotes the weight of Home goods in the Foreign aggregate consumption basket).

Asset Portfolio and Budget Constraint

Agents have access to country-specific equity, short and long-term bonds. All assets are traded on globally integrated financial markets, and are denominated in the currency of the issuing country. Each asset is assumed to be in zero net supply.

Let $B_{jk,t}$ ($B_{jk,t}^L$) denote the stock of real external holdings of short-term (long-term) bonds,

¹⁶For $a_H > 1/2$, there is ‘home bias’ in consumption.

issued by country k , held by an agent in country j , carried from period t to $t+1$, defined in units of the Home consumption basket, with $j, k = \{H, F\}$.¹⁷ Similarly, let $S_{jk,t}$ denote the stock of real external holdings of country- k equity held by an agent in country j , carried from period t to $t+1$, also defined in terms of the Home consumption basket. Because $B_{HH,t}$, $B_{HH,t}^L$ and $S_{HH,t}$ are defined as the ‘external’ holdings of Home assets by Home households, then the concurrent Foreign holdings of Home assets are $B_{FH,t} = -B_{HH,t}$, $B_{FH,t}^L = -B_{HH,t}^L$ and $S_{FH,t} = -S_{HH,t}$.

The Home household flow budget constraint, in units of the Home consumption basket, is:

$$\begin{aligned} C_t + B_{HH,t} + e^{-\zeta_{F,t}} B_{HF,t} + B_{HH,t}^L + e^{-\zeta_{F,t}} B_{HF,t}^L + \\ e^{-\zeta_{S,t}} S_{HH,t} + e^{-\zeta_{F,t}} e^{-\zeta_{S^*,t}} S_{HF,t} = w_t L_t + \Pi_t - T_t + B_{HH,t-1} r_t + \\ B_{HF,t-1} r_t^* + B_{HH,t-1}^L r_{L,t} + B_{HF,t-1}^L r_{L,t}^* + S_{HH,t-1} r_{e,t} + S_{HF,t-1} r_{e,t}^* \end{aligned} \quad (5.8)$$

where w_t denotes the Home real wage, Π_t denotes the real profits of Home firms paid as dividends to the Home households, and T_t denotes real lump-sum taxes levied by the government.¹⁸ r_t (r_t^*) represents the gross real return on Home (Foreign) short-term bonds purchased in period $t-1$; $r_{L,t}$ ($r_{L,t}^*$) is the gross real one-period return on Home (Foreign) long-term bonds purchased in period $t-1$; and $r_{e,t}$ ($r_{e,t}^*$) denotes the gross real return on Home (Foreign) equity purchased in period $t-1$. All gross real returns are defined in units of the Home consumption basket.

There are three shocks in the Home budget constraint: $\zeta_{F,t}$, $\zeta_{S,t}$ and $\zeta_{S^*,t}$. They influence the relative demand for assets and are uncorrelated with one another. $\zeta_{F,t}$ is an (inverse) cost shock to the ability of households to trade Foreign assets. Itskhoki and Mukhin (2017) label this an ‘international asset demand shock’,¹⁹ influencing the relative demand for Home versus Foreign assets. The international asset demand shock process is:

$$\zeta_{F,t} = \rho_F \zeta_{F,t-1} + \varepsilon_{F,t}, \quad \varepsilon_{F,t} \sim \mathcal{N}(0, \sigma_F^2), \quad \rho_F \in (0, 1), \quad \sigma_F > 0 \quad (5.9)$$

$\zeta_{S,t}$ ($\zeta_{S^*,t}$) is a shock to the demand for Home (Foreign) equity. It can be interpreted as a reduced form characterisation of an external finance premium shock, generating exogenous variation in equity returns. The equity demand shock process is:

$$\zeta_{S,t} = \rho_S \zeta_{S,t-1} + \varepsilon_{S,t}, \quad \varepsilon_{S,t} \sim \mathcal{N}(0, \sigma_S^2), \quad \rho_S \in (0, 1), \quad \sigma_S > 0 \quad (5.10)$$

and similarly for the Foreign equity demand shock, with $\rho_S = \rho_S^*$ and $\sigma_S = \sigma_S^*$.

¹⁷Despite bonds being nominal, I mathematically define the stock of bond holdings in real terms for computational convenience. This definition lends itself to the application of model solution techniques proposed by Devereux and Sutherland (2010, 2011) for international macroeconomic models with portfolios of international assets, as outlined in section 5.3.6.

¹⁸Here, the Home agent receives all Home profits through dividends in the first instance, while claims to those profits are traded with equity. In a symmetric equilibrium with zero net foreign assets, gross portfolio holdings exactly offset each other in value terms. That is, if $S_{HH,t} < 0$ (implying $S_{FH,t} > 0$ by definition of external asset holdings), then Foreign households hold some non-negative claim to Home profits. As in Devereux and Sutherland (2008), this is simply an accounting convention which simplifies the exposition of the model, but is not critical. Alternatively, one can treat all dividend income as part of equity returns, so that wage earnings represent the Home agents’ only non-financial income. In this case, even in a symmetric equilibrium with zero net foreign assets, agents in each economy would have non-zero *net* portfolio positions.

¹⁹Itskhoki and Mukhin (2017) argue that this shock has a variety of interpretations, which are all isomorphic in reduced form, including that the shock: limits arbitrage in currency markets; represents heterogeneous beliefs in currency markets; or represents financial frictions in currency markets (Gabaix and Maggiori, 2015). In all cases the shock admits deviations from uncovered interest parity (UIP).

Asset Returns and Exchange Rates

Home equity represents a claim on the profits of Home firms. Let $Z_{E,t}$ represent the real price of Home equity at time t in units of the Home consumption basket. Then, the gross real rate of return on Home equity is:

$$r_{e,t} = \frac{\Pi_t + Z_{E,t}}{Z_{E,t-1}}$$

The real price of Foreign equity at time t in units of the Foreign consumption basket is $Z_{E,t}^*$. In units of the Home consumption basket this price is $Q_t Z_{E,t}^*$, where Q_t is the real exchange rate defined as:

$$Q_t = \frac{\mathcal{E}_t P_t^*}{P_t} \quad (5.11)$$

and \mathcal{E}_t is the nominal exchange rate — the Home country price of one unit of Foreign currency — defined such that an increase in \mathcal{E}_t (Q_t) represents a nominal (real) depreciation of Home currency. The gross real rate of return on Foreign equity in units of the Home consumption basket can be written as:

$$r_{e,t}^* = \frac{Q_t}{Q_{t-1}} \frac{\Pi_t^* + Z_{E,t}^*}{Z_{E,t-1}^*}$$

Home nominal short-term bonds represent a claim to one unit of Home currency in the subsequent period. The real price, in units of the Home consumption basket, of a Home short-term nominal bond purchased in period $t-1$ is Z_{t-1} . The real payoff of this bond when carried into period t is $1/P_t$. Thus, the gross real rate of return on the Home nominal bond is $r_t = \frac{1}{P_t Z_{t-1}}$ in units of the Home consumption basket. In units of Home currency, the gross nominal rate of return on the Home nominal bond is $R_{t-1} = r_t \pi_t$, where $\pi_t \equiv P_t/P_{t-1}$ and R_t denotes the nominal short-term interest rate from time t to $t+1$, such that $R_{t-1} = \frac{1}{P_{t-1} Z_{t-1}}$.

Foreign nominal short-term bonds represent a claim to one unit of Foreign currency in the subsequent period. The real price, in units of the Foreign (Home) consumption basket, of a Foreign short-term nominal bond purchased in period $t-1$ is Z_{t-1}^* ($Q_{t-1} Z_{t-1}^*$). The gross real rate of return on the Foreign nominal bond is $r_t^* = \frac{Q_t}{Q_{t-1} P_t^* Z_{t-1}^*}$ in units of the Home consumption basket. In units of Foreign currency, the gross nominal rate of return on the Foreign nominal bond is $R_{t-1}^* = r_t^* \pi_t \frac{\mathcal{E}_{t-1}}{\mathcal{E}_t} = \frac{1}{P_{t-1}^* Z_{t-1}^*}$.

Following Woodford (2001), long-term bonds are modelled as perpetuities. The period- t nominal price of a Home long-term bond, newly-issued in period t , is $P_{L,t}$, in units of the issuing country's currency. Thereafter, the bond pays an exponentially decaying nominal coupon κ^s at time $t+s+1$ for $s = 0, 1, 2, \dots$ and $\kappa \in (0, 1]$, also expressed in units of the issuing country's currency.²⁰ That is, the Home long-term bond pays 1 unit of Home currency in period $t+1$, κ units in $t+2$, κ^2 units in $t+3$, etc. The one-period gross nominal yield to maturity at time t on this bond is $R_{L,t} = \frac{1}{P_{L,t}} + \kappa$.²¹ The gross one-period real rate of return $r_{L,t}$ on the long-term Home bond, in units of the Home consumption basket, is related to the nominal yield

²⁰If $\kappa = 1$, the security is a consol.

²¹The price of a nominal bond is equal to the present discounted value of its future payments, so when $\kappa < R_{L,t}$:

$$P_{L,t} = \frac{1}{R_{L,t}} + \frac{\kappa}{R_{L,t}^2} + \frac{\kappa^2}{R_{L,t}^3} + \dots = \frac{1}{R_{L,t}} \frac{1}{1 - \frac{\kappa}{R_{L,t}}} = \frac{1}{R_{L,t} - \kappa}$$

to maturity through the following expression:²²

$$r_{L,t} = \frac{1}{\pi_t} \frac{P_{L,t}}{P_{L,t-1}} R_{L,t}$$

The Foreign long-term bond, for which the nominal price $P_{L,t}^*$ and coupon κ^* are expressed in Foreign currency units, has nominal yield to maturity at time t of $R_{L,t}^* = \frac{1}{P_{L,t}^*} + \kappa^*$ in units of Foreign currency. The gross one-period real rate of return, in units of the Home consumption basket, is related to this through the following expression:

$$r_{L,t}^* = \frac{1}{\pi_t} \frac{\mathcal{E}_t}{\mathcal{E}_{t-1}} \frac{P_{L,t}^*}{P_{L,t-1}^*} R_{L,t}^*$$

Household Optimality Conditions

The optimality conditions for the Home households are:

$$w_t = e^{\zeta_{L,t}} C_{X,t}^\sigma L_t^{\frac{1}{\nu}} \quad (5.12)$$

$$1 = \mathbb{E}_t \left[\beta(C_t) e^{\Delta \zeta_{C,t+1}} e^{\zeta_{i,t}} \left(\frac{C_{X,t+1}}{C_{X,t}} \right)^{-\sigma} r_{i,t+1} \right], \quad \text{where } i = 1, 2, \dots, 6 \quad (5.13)$$

where $C_{X,t} \equiv C_t - \gamma C_{t-1}$, $\zeta_{1,t} \equiv 0$, $\zeta_{2,t} \equiv \zeta_{F,t}$, $\zeta_{3,t} \equiv 0$, $\zeta_{4,t} \equiv \zeta_{F,t}$, $\zeta_{5,t} \equiv \zeta_{S,t}$, $\zeta_{6,t} \equiv \zeta_{F,t} + \zeta_{S^*,t}$, $r_{1,t+1} \equiv r_{t+1}$, $r_{2,t+1} \equiv r_{t+1}^*$, $r_{3,t+1} \equiv r_{L,t+1}$, $r_{4,t+1} \equiv r_{L,t+1}^*$, $r_{5,t+1} \equiv r_{e,t+1}$ and $r_{6,t+1} \equiv r_{e,t+1}^*$.²³ Equation (5.12) is the Home intratemporal Euler equation associated with optimal labour supply. Equation (5.13) represents the Home intertemporal Euler equations for the Home and Foreign assets; they comprise the portfolio optimality conditions necessary for the application of the Devereux and Sutherland (2010, 2011) solution method. Equivalent optimality conditions for the Foreign household are presented in appendix D.1.1.

5.3.2 Firms

Output in each country is produced by a continuum of monopolistically competitive firms. The production function for a good produced by firm $h \in (0, 1)$ in the Home country is:

$$Y_t(h) = e^{a_t} L_t(h)^{1-\alpha} X_t(h)^\alpha \quad (5.14)$$

where $L_t(h)$ is the labour input for firm h , $\alpha \in (0, 1)$ is the elasticity of output with respect to intermediate goods $X_t(h)$, which are the same bundle of Home and Foreign goods as in household consumption, with price index P_t in the Home economy.²⁴ a_t is a stochastic productivity shock with evolution:

$$a_t = \rho_a a_{t-1} + \varepsilon_{a,t}, \quad \varepsilon_{a,t} \sim \mathcal{N}(0, \sigma_a^2), \quad \rho_a \in (0, 1), \quad \sigma_a > 0 \quad (5.15)$$

and similarly for the Foreign productivity shock, with $\rho_a = \rho_a^*$ and $\sigma_a = \sigma_a^*$. Motivated by evidence in Benigno and Thoenissen (2008), I allow the Home and Foreign productivity shocks

²²Appendix D.1.1 provides a derivation of this expression by linking the Home household budget constraint in units of the Home consumption basket (5.8) to a nominal equivalent, expressed in units of the Home currency.

²³The numerical labels for $\zeta_{i,t}$ and $r_{i,t}$ are adopted to simplify the algebraic exposition in section 5.3.6.

²⁴The corresponding production function for a Foreign firm $f \in (0, 1)$ is $Y_t^*(f) = e^{a_t^*} L_t^*(f)^{1-\alpha} X_t^*(f)^\alpha$, where $X_t^*(f)$ represents intermediate goods with price index P_t^* .

to be correlated with one another, but assume they are uncorrelated with all other shocks.

The Home real marginal cost of production is:²⁵

$$mc_t = e^{-a_t} \frac{w_t^{1-\alpha}}{(1-\alpha)^{1-\alpha} \alpha^\alpha} \quad (5.16)$$

Because all firms face the same real marginal costs, optimal Home input demands are:

$$L_t = \frac{(1-\alpha)mc_t Y_t}{w_t} \quad (5.17)$$

$$X_t = \alpha mc_t Y_t \quad (5.18)$$

where (5.17) is labour demand and (5.18) is intermediate goods demand.

Home firms sell their produce to consumers, firms and government in the Home and Foreign economies. The nominal profits of the Home firm can therefore be written as:

$$P_t \Pi_t = (P_{H,t} - P_t mc_t) Y_{H,t} + (\mathcal{E}_t P_{H,t}^* - P_t mc_t) Y_{H,t}^* \quad (5.19)$$

where $Y_{H,t}$ denotes Home demand for the basket of Home goods and $Y_{H,t}^*$ denotes Foreign demand for the basket of Home goods. Demand for intermediate goods (5.18), along with the expenditure minimisation problem associated with (5.4) for the Home household and a similar problem for the Foreign household, implies that these quantities can be written as:

$$Y_{H,t} = a_H \left(\frac{P_{H,t}}{P_t} \right)^{-\phi} (C_t + \alpha mc_t Y_t + G_t) \quad (5.20)$$

$$Y_{F,t} = (1 - a_H) \left(\frac{P_{F,t}}{P_t} \right)^{-\phi} (C_t + \alpha mc_t Y_t + G_t) \quad (5.21)$$

where G_t denotes Home government spending, which is the same bundle of Home and Foreign goods as in Home household consumption (5.4). Home demand for the basket of Foreign goods $Y_{F,t}$ and Foreign demand for the basket of Foreign goods $Y_{F,t}^*$ are given by:

$$Y_{H,t}^* = a_H^* \left(\frac{P_{H,t}^*}{P_t^*} \right)^{-\phi} (C_t^* + \alpha mc_t^* Y_t^* + G_t^*) \quad (5.22)$$

$$Y_{F,t}^* = (1 - a_H^*) \left(\frac{P_{F,t}^*}{P_t^*} \right)^{-\phi} (C_t^* + \alpha mc_t^* Y_t^* + G_t^*) \quad (5.23)$$

where G_t^* denotes Foreign government spending, which is also the same bundle of Home and Foreign goods as in Foreign household consumption.

Pricing

The monopolistically competitive firms have price-setting power, subject to nominal rigidities à la Calvo (1983), such that, at any time t , firms are unable to change their price with fixed probability $\xi = \xi^* \in [0, 1)$. I assume that firms updating their price do so simultaneously in Home and Foreign markets. Firms face producer currency pricing (PCP), setting all prices in

²⁵See appendix D.1.2 for a derivation of (5.16)-(5.18).

their domestic currency (as in, e.g. Obstfeld and Rogoff, 1995).²⁶

Let $\mathcal{P}_t(h)$ denote the price of the Home good in the Home market optimally chosen by firm h who resets its price at time t . $\mathcal{E}_t \mathcal{P}_t^*(h)$ denotes the price set by firm h for the Foreign market, in Home currency terms. Under PCP, \mathcal{E}_t and $\mathcal{P}_t^*(h)$, move inversely — i.e. there is complete exchange rate pass through — and the pricing problem of Home firms is:

$$\max_{\{\mathcal{P}_t(h), \mathcal{E}_t \mathcal{P}_t^*(h)\}} \mathbb{E}_t \sum_{s=0}^{\infty} \delta_s \xi^s \Omega_{t+s} \left[\frac{(\mathcal{P}_t(h))^{1-\theta}}{(P_{H,t+s})^{-\theta}} Y_{H,t+s} + \frac{(\mathcal{E}_t \mathcal{P}_t^*(h))^{1-\theta}}{(\mathcal{E}_{t+s} P_{H,t+s}^*)^{-\theta}} Y_{H,t+s}^* \right. \\ \left. - P_{t+s} m c_{t+s} \left(\frac{\mathcal{P}_t(h)}{P_{H,t+s}} \right)^{-\theta} Y_{H,t+s} - P_{t+s} m c_{t+s} \left(\frac{\mathcal{E}_t \mathcal{P}_t^*(h)}{\mathcal{E}_{t+s} P_{H,t+s}^*} \right)^{-\theta} Y_{H,t+s}^* \right] \quad (5.24)$$

where $\Omega_{t+s} \equiv \frac{u'(C_{t+s})}{u'(C_t)}$ is the discount factor used to evaluate Home firm profits. The first two terms in square brackets represent Home firm revenues from sales to Home and Foreign, while the final two terms express the costs of producing this output.

The Home firm's problem is solved by the following optimality conditions:

$$\mathcal{P}_t(h) = e^{\mu_t} \frac{\theta}{\theta - 1} \frac{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s \xi^s \Omega_{t+s} m c_{t+s} P_{t+s} P_{H,t+s}^{\theta} Y_{H,t+s}}{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s \xi^s \Omega_{t+s} P_{H,t+s}^{\theta} Y_{H,t+s}} \\ \mathcal{E}_t \mathcal{P}_t^*(h) = e^{\mu_t} \frac{\theta}{\theta - 1} \frac{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s \xi^s \Omega_{t+s} m c_{t+s} P_{t+s} (\mathcal{E}_{t+s} P_{H,t+s}^*)^{\theta} Y_{H,t+s}^*}{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s \xi^s \Omega_{t+s} (\mathcal{E}_{t+s} P_{H,t+s}^*)^{\theta} Y_{H,t+s}^*} \quad (5.25)$$

where μ_t has been added to represent a Home firm markup shock, with evolution:

$$\mu_t = \rho_{\mu} \mu_{t-1} + \varepsilon_{\mu,t}, \quad \varepsilon_{\mu,t} \sim \mathcal{N}(0, \sigma_{\mu}^2), \quad \rho_{\mu} \in (0, 1), \quad \sigma_{\mu} > 0 \quad (5.26)$$

Symmetric pricing expressions exist for Foreign firms, with Foreign markup shock μ_t^* , where $\rho_{\mu} = \rho_{\mu}^*$ and $\sigma_{\mu} = \sigma_{\mu}^*$.

Because all producers that reset their price in period t optimally choose the same price, the dynamic evolution of $P_{H,t}$ and $P_{F,t}^*$ can be written as:²⁷

$$P_{H,t}^{1-\theta} = \xi P_{H,t-1}^{1-\theta} + (1 - \xi) \mathcal{P}_t(h)^{1-\theta} \quad (5.27)$$

$$(P_{F,t}^*)^{1-\theta} = \xi (P_{F,t-1}^*)^{1-\theta} + (1 - \xi) (\mathcal{P}_t^*(f))^{1-\theta} \quad (5.28)$$

5.3.3 Monetary Policy

Monetary policy in the Home and Foreign economies follow Taylor-type rules, with policymakers setting the short-term nominal interest rate in their domestic currency. The interest rate rule for the Home economy is:

$$R_t = R_{t-1}^{\rho_r} \left[\bar{\beta}^{-1} \left(\frac{P_t}{P_{t-1}} \right)^{\phi_{\pi}} \left(\frac{Y_t}{\bar{Y}} \right)^{\phi_y} \right]^{1-\rho_r} e^{\varepsilon_{mp,t}} \quad (5.29)$$

²⁶In appendix D.2, I report results under two alternative pricing models: (i) local currency pricing (LCP) (e.g. Betts and Devereux, 2000), where firms set prices in the currency of the market in which they sell the goods; and (ii) dollar currency pricing (DCP), where both the Home country and the US (the Foreign country) invoice their export prices in dollar terms (i.e. in the foreign currency), motivated by evidence in Gopinath (2015) that the dollar is the dominant currency in world trade.

²⁷Appendix D.1.2 presents a derivation.

where $\rho_r \in [0, 1]$ denotes the degree of interest rate smoothing, $\phi_\pi > 1$ and $\phi_y > 0$. \bar{Y} represents steady state output of the Home country. $\varepsilon_{mp,t}$ is a stochastic monetary policy disturbance with distribution: $\varepsilon_{mp,t} \sim \mathcal{N}(0, \sigma_{mp}^2)$, where $\sigma_{mp} > 0$. For simplicity, there is a symmetric interest rate rule for the Foreign economy, with $\rho_r = \rho_r^*$, $\phi_\pi = \phi_\pi^*$, $\phi_y = \phi_y^*$, $\sigma_{mp} = \sigma_{mp}^*$.

This form of interest rate rule is chosen to replicate the actual practice of central banks, as opposed to an optimal rule from a welfare perspective (e.g. Corsetti et al., 2010, 2016). Specifically, the nominal interest rate is a function of CPI inflation within a country, as has been the case in countries following inflation targeting policies in recent decades.

5.3.4 Government

Government spending is exogenous and subject to stochastic shocks, evolving according to:

$$G_t = \rho_G G_{t-1} + \varepsilon_{G,t}, \quad \varepsilon_{G,t} \sim \mathcal{N}(0, \sigma_G^2), \quad \rho_G \in (0, 1), \quad \sigma_G > 0 \quad (5.30)$$

with a symmetric, and independent, expression for Foreign government spending G_t^* , with $\rho_G = \rho_G^*$, $\sigma_G = \sigma_G^*$. All government spending is financed by the lump-sum taxes levied on households T_t , so that the government budget constraint is: $G_t = T_t$.²⁸

5.3.5 Equilibrium

In equilibrium, the product, labour and asset markets must clear.

The Home labour market clears when L_t is consistent with labour supply (5.12) and labour demand (5.17), and symmetrically for L_t^* in the Foreign country.

Goods market clearing in Home and Foreign markets requires that aggregate output is equal to the sum of demand from domestic and foreign sources:

$$Y_t = Y_{H,t} + Y_{H,t}^* \quad (5.31)$$

$$Y_t^* = Y_{F,t} + Y_{F,t}^* \quad (5.32)$$

where $Y_{H,t}$, $Y_{F,t}$, $Y_{H,t}^*$ and $Y_{F,t}^*$ are consistent with (5.20)-(5.23).

Asset market clearing requires that all assets are in zero net supply internationally, so:

$$\alpha_{i,t} + \alpha_{i,t}^* = 0, \quad \text{where } i = 1, 2, \dots, 6 \quad (5.33)$$

where $\alpha_{1,t} \equiv B_{HH,t}$, $\alpha_{2,t} \equiv B_{HF,t}$, $\alpha_{3,t} \equiv B_{HH,t}^L$, $\alpha_{4,t} \equiv B_{HF,t}^L$, $\alpha_{5,t} \equiv S_{HH,t}$, $\alpha_{6,t} \equiv S_{HF,t}$, $\alpha_{1,t}^* \equiv B_{FH,t}$, $\alpha_{2,t}^* \equiv B_{FF,t}$, $\alpha_{3,t}^* \equiv B_{FH,t}^L$, $\alpha_{4,t}^* \equiv B_{FF,t}^L$, $\alpha_{5,t}^* \equiv S_{FH,t}$ and $\alpha_{6,t}^* \equiv S_{FF,t}$.

Finally, the steady state Home household budget constraint (5.8) can be written as:

$$P_t \sum_{i=1}^6 \alpha_{i,t} - P_t \sum_{i=1}^6 \alpha_{i,t-1} r_{i,t} = \mathcal{E}_t P_{H,t}^* Y_{H,t}^* - P_{F,t} Y_{F,t} \quad (5.34)$$

by setting shocks to their steady state value, and substituting the expressions for Home nominal profits (5.19), Home labour demand (5.17), and demand equations (5.20)-(5.23) into (5.19).²⁹

²⁸ $G_t < 0$ can be interpreted as a lump-sum transfer to households.

²⁹Appendix D.1.3 provides a derivation of this expression.

5.3.6 Model Solution

A standard first-order approximation of the model will not yield solutions to the portfolio problem for two reasons. First, the first-order model approximation is not sufficient to determine the optimal portfolio. Second, in the non-stochastic steady state, the equilibrium portfolio is indeterminate.³⁰ To overcome this, I use the method of Devereux and Sutherland (2011) to solve for the optimal zero-order (or steady state) portfolio holdings, and its extension in Devereux and Sutherland (2010) to solve for the optimal time-varying asset portfolio allocation in response to shocks.³¹ Devereux and Sutherland (2011) show that the first of these problems can be overcome by using higher-order approximations of the portfolio problem, and the second can be overcome by treating the approximation point — the steady state portfolio — as unknown.

The solution method for steady state portfolio holdings relies on two observations. First, that the first-order behaviour of macroeconomic variables is independent of time-variation in the portfolio allocation; only the steady state value of the portfolio allocation influences the first-order behaviour of macroeconomic variables. Second, that the steady state value of the portfolio allocation can be attained by combining a second-order approximation of the portfolio equations (5.13) with a first-order approximation of the non-portfolio equations.

In order to solve for the dynamic behaviour of asset holdings around the steady state portfolio, higher-order approximations are required. Devereux and Sutherland (2010) show that a third-order approximation of the portfolio equations captures the first-order effect of state variables on the second moments of asset returns. Thus, a third-order approximation of the model portfolio equations can be used in conjunction with first and second-order approximations of the non-portfolio equations to solve for dynamic of optimal portfolios.

Hereafter, a bar over a variable indicates its steady state value \bar{x} and a hat indicates the log-deviation from the steady state $\hat{x}_t \equiv \ln(x_t/\bar{x})$, unless otherwise stated.

Portfolio Equations

To apply the Devereux and Sutherland (2010, 2011) solution method, the steady state Home household budget constraint (5.34) can be rewritten as:

$$P_t NFA_t = P_t \left[NFA_{t-1} r_{1,t} + \sum_{i=2}^6 \alpha_{i,t-1} (r_{i,t} - r_{1,t}) \right] + \mathcal{E}_t P_{H,t}^* Y_{H,t}^* - P_{F,t} Y_{F,t} \quad (5.35)$$

where the net foreign assets of Home agents NFA_t have the following definition using (5.33):³²

$$NFA_t \equiv \sum_{i=1}^6 \alpha_{i,t} \quad (5.36)$$

³⁰Devereux and Sutherland (2010) emphasise that these are two distinct problems. The first arises in the approximated model with stochastic shocks because certainty equivalence holds, while the second arises in the non-approximated model without stochastic shocks.

³¹Because this model includes multiple nominal assets, this application of the solution method is most similar to Devereux and Sutherland (2007, 2008). Devereux and Sutherland (2010, 2011) outline the solution method in real models, while Devereux and Sutherland (2007, 2008) apply the method in nominal models.

³²For example, net holdings of short-term bonds in the Home country can be rewritten using (5.33):

$$B_{HF,t} - B_{FH,t} = \alpha_{2,t} - \alpha_{1,t}^* = \alpha_{1,t} + \alpha_{2,t}$$

At the end of each period, agents from both countries select an asset portfolio to carry into the following period. This is optimally determined by the households' intertemporal Euler equations. Let the Home short-term bond $\alpha_{1,t}$ act as the numéraire asset. The six Home intertemporal Euler equations (5.13) can be reduced to five portfolio optimality conditions:

$$\mathbb{E}_t \left[C_{X,t+1}^{-\sigma} \tilde{r}_{1,t+1} \right] = \mathbb{E}_t \left[C_{X,t+1}^{-\sigma} \tilde{r}_{i,t+1} \right], \quad \text{where } i = 2, 3, \dots, 6 \quad (5.37)$$

where $\tilde{r}_{i,t+1} \equiv e^{\Delta \zeta_{C,t+1}} e^{\zeta_{i,t}} r_{i,t+1}$ for all $i = 1, 2, \dots, 6$. Equivalent Foreign expressions are:

$$\mathbb{E}_t \left[\mathcal{Q}_{t+1}^{-1} (C_{X,t+1}^*)^{-\sigma} \tilde{r}_{1,t+1} \right] = \mathbb{E}_t \left[\mathcal{Q}_{t+1}^{-1} (C_{X,t+1}^*)^{-\sigma} \tilde{r}_{i,t+1} \right], \quad \text{where } i = 2, 3, \dots, 6 \quad (5.38)$$

The two sets of portfolio optimality conditions, (5.37) and (5.38), and the market clearing conditions (5.33) for $i = 2, 3, \dots, 6$, provide fifteen equations which determine $\alpha_{i,t}$, $\alpha_{i,t}^*$ and $r_{x,i,t+1} \equiv r_{i,t+1} - r_{1,t+1}$ for $i = 2, 3, \dots, 6$.

Steady State Portfolio Allocation

Given the scale of this model, I numerically solve for the steady state asset portfolio by applying the Devereux and Sutherland (2011) method, and discuss the findings in section 5.4. In this subsection I outline the solution method.

I approximate the non-portfolio equations around the non-stochastic, zero-growth, zero-inflation steady state of the model. I assume that both countries are symmetric, so countries have zero net external assets in steady state — $\overline{NFA} = 0$, $\overline{Q} = \overline{\mathcal{E}} = 1$ — and all other macroeconomic variables are equal across countries — e.g. $\overline{Y} = \overline{Y}^*$, $\overline{C}_X = \overline{C}_X^*$ etc.

A second-order approximation of the Home portfolio equations (5.37), which is required to solve for the steady state optimal asset portfolio allocation, yields:

$$\mathbb{E}_t \left[\hat{r}_{x,i,t+1} + \frac{1}{2} \left(\hat{r}_{i,t+1}^2 - \hat{r}_{1,t+1}^2 \right) - \sigma \hat{C}_{X,t+1} \hat{r}_{x,i,t+1} \right] = 0 + \mathcal{O}(\epsilon^3), \quad \text{where } i = 2, 3, \dots, 6$$

where $\hat{r}_{x,i,t+1} \equiv \hat{r}_{i,t+1} - \hat{r}_{1,t+1}$ and $\mathcal{O}(\epsilon^3)$ is the residual containing all terms of order higher than two. These second-order approximations can be stacked into vector form:

$$\mathbb{E}_t \left[\hat{\mathbf{r}}_{x,t+1} + \frac{1}{2} \hat{\mathbf{r}}_{x,t+1}^2 - \sigma \hat{C}_{X,t+1} \hat{\mathbf{r}}_{x,t+1} \right] = 0 + \mathcal{O}(\epsilon^3) \quad (5.39)$$

where $\hat{\mathbf{r}}_{x,t+1} \equiv \left[\hat{r}_{x,2,t+1}, \hat{r}_{x,3,t+1}, \dots, \hat{r}_{x,6,t+1} \right]'$ and

$$\hat{\mathbf{r}}_{x,t+1}^2 \equiv \left[\hat{r}_{2,t+1}^2 - \hat{r}_{1,t+1}^2, \hat{r}_{3,t+1}^2 - \hat{r}_{1,t+1}^2, \dots, \hat{r}_{6,t+1}^2 - \hat{r}_{1,t+1}^2 \right]'$$

The corresponding expression for the Foreign portfolio conditions (5.38) is:

$$\mathbb{E}_t \left[\hat{\mathbf{r}}_{x,t+1} + \frac{1}{2} \hat{\mathbf{r}}_{x,t+1}^2 - \sigma \hat{C}_{X,t+1}^* \hat{\mathbf{r}}_{x,t+1} - \hat{\mathcal{Q}}_{t+1} \hat{\mathbf{r}}_{x,t+1} \right] = 0 + \mathcal{O}(\epsilon^3) \quad (5.40)$$

(5.39) and (5.40) can be combined to yield two equilibrium conditions. First, their difference:

$$\mathbb{E}_t \left[\left(\hat{C}_{X,t+1} - \hat{C}_{X,t+1}^* - \hat{\mathcal{Q}}_{t+1}/\sigma \right) \hat{\mathbf{r}}_{x,t+1} \right] = 0 + \mathcal{O}(\epsilon^3) \quad (5.41)$$

and second, their sum:

$$\mathbb{E}_t \left[\hat{\mathbf{r}}_{x,t+1} \right] = -\frac{1}{2} \mathbb{E}_t \left[\hat{\mathbf{r}}_{x,t+1}^2 \right] + \sigma \frac{1}{2} \mathbb{E}_t \left[\left(\hat{C}_{X,t+1} + \hat{C}_{X,t+1}^* + \hat{Q}_{t+1}/\sigma \right) \hat{\mathbf{r}}_{x,t+1} \right] + \mathcal{O}(\epsilon^3) \quad (5.42)$$

(5.41) provides a sufficient condition for pinning down steady state values of portfolio holdings.

(5.42) provides corresponding conditions for equilibrium expected excess returns.

Devereux and Sutherland (2011) highlight three key properties of the solution method. First, to evaluate the left-hand side of (5.41), it is sufficient to derive expressions for the first-order behaviour of consumption and excess returns — i.e. first-order approximations of non-portfolio equations. Second, only the zero-order portfolio allocation affects the first-order accurate behaviour of consumption and excess returns; higher-order aspects of the portfolio decision are irrelevant for the first-order accurate behaviour of non-portfolio variables.³³ Third, to a first-order approximation, the portfolio excess return is a zero mean i.i.d. random variable. This follows from (5.42) because it only contains second-order terms. This observation simplifies the solution method, as it implies that the steady state portfolio does not affect the eigenvalues of the first-order macroeconomic system.

Portfolio Dynamics

In section 5.5, I go beyond the steady state portfolio holdings and study the response of interest rates and portfolio positions to macroeconomic shocks. I use the procedure proposed by Devereux and Sutherland (2010) to account for optimal portfolio dynamics.

Devereux and Sutherland (2010) show that, to solve for the dynamic behaviour of asset holdings around the steady state portfolio, it is necessary to carry out a third-order approximation of the portfolio equations. This provides information about how changes in state variables influence the risk characteristics of assets. The third-order approximation of the portfolio equations can be combined with first and second-order approximations of the non-portfolio equations to solve for the optimal portfolio dynamics. Devereux and Sutherland (2010) derive analytical expressions for the optimal portfolio dynamics within their model. However, because of the scale of this model, I derive the portfolio dynamics numerically by applying their algorithm.

5.4 International Risk Sharing

In this section, I present the model calibration, describe its ability to replicate the stylised facts presented in section 5.2, and discuss the mechanisms underlying international risk sharing.

5.4.1 Benchmark Calibration

This quarterly frequency benchmark model calibration is listed in table 5.2.

Structural Parameters The steady state value of the discount factor $\bar{\beta}$ is chosen to yield a steady state annualised (real and nominal) rate of return of approximately 4%. The Uzawa convergence parameter η is set such that the speed of convergence to the nonstochastic steady state is small. The constant term in the endogenised discount factor ω (ω^*) is implicitly defined given $\bar{\beta}$, \bar{C} and η .

³³This is the case because steady state asset returns are equal.

Table 5.2: Benchmark Parameter Values

PARAMETER	DESCRIPTION	VALUE
<u>STRUCTURAL PARAMETERS</u>		
$\bar{\beta} = \omega \bar{C}^{-\eta}$	Steady State Discount Factor	0.99
η	Uzawa Convergence Parameter	0.01
σ	Coefficient of Relative Risk Aversion	2
γ	Persistence of Habit Stock	0.65
ν	Frisch Labour Supply Elasticity	0.5
a_H	Share of Home Goods in Consumption Basket	0.72
θ	Brand Elasticity	7.7
κ (κ^*)	Home (Foreign) Nominal Coupon on Long-Term Bond	0.9595
$1 - \alpha$	Labour Share in Production of Final Goods	0.61
ξ	Fraction of Prices not Reset Each Period	0.3023
ρ_r	Degree of Monetary Policy Smoothing	0.91
ϕ_π	Taylor Rule Reaction to CPI Inflation	1.58
ϕ_y	Taylor Rule Reaction to Output Deviations from Steady State	0.14
ϕ	Trade Elasticity	0.78
<u>SHOCK PROCESS PARAMETERS</u>		
ρ_C	Persistence of Preference Shock	0.74
σ_C	Standard Deviation of Preference Shock	0.0040
ρ_L	Persistence of Labour Supply Shock	0.88
σ_L	Standard Deviation of Labour Supply Shock	0.0042
ρ_a	Persistence of Productivity Shock	0.95
σ_a	Standard Deviation of Productivity Shock	0.0134
$\text{Cov}(a, a^*)$	Covariance of Home and Foreign Productivity Shocks	0.000081
ρ_F	Persistence of International Asset Demand Shock	0.97
σ_F	Standard Deviation of International Asset Demand Shock	0.0005
ρ_S	Persistence of Equity Demand Shock	0
σ_S	Standard Deviation of Equity Demand Shock	0.0049
ρ_μ	Persistence of Markup Shock	0
σ_μ	Standard Deviation of Markup Shock	0.0015
σ_{mp}	Standard Deviation of Monetary Policy Shock	0.0013
ρ_G	Persistence of Government Spending Shock	0.9
σ_G	Standard Deviation of Government Spending Shock	0.0030

Note: Parameter calibration for the model presented in section 5.3. The model is calibrated at a quarterly frequency.

The coefficient of relative risk aversion σ is set to 2 as in existing studies into international capital flows (e.g. Corsetti et al., 2008b). The calibration for habit persistence γ is motivated by Smets and Wouters (2007). The Frisch elasticity of labour supply ν is chosen to match the figure reported in Chetty, Guren, Manoli, and Weber (2011) for the intensive margin.

The relative weight on Home tradables in total tradables consumption a_H is chosen such that imports are 5% of aggregate output in steady state. Corsetti et al. (2008b) choose this figure as it corresponds to the average ratio of US imports from Europe, Canada and Japan to US GDP between 1960 and 2002. The elasticity of substitution between varieties of tradable goods θ — the brand elasticity — is chosen to yield a steady state monopoly markup of 15%, as in Corsetti et al. (2008a).

The long-term bond parameters, κ and κ^* , denote the nominal coupon in units of the issuing country's currency. Woodford (2001) shows that these perpetuities have duration $(1 - \bar{\beta}\kappa)^{-1}$ when prices are stable. The parameters are chosen to imply a duration of 5 years, approximately in line with estimates of the duration of outstanding privately-held debt reported in Hilscher, Raviv, and Reis (2014).

The labour share in the production of final tradable goods $1 - \alpha$ is set to match Stockman and Tesar (1995), who calculate that labour comprises 61% of the production share of tradable goods. The Calvo-pricing parameter ξ is chosen to imply an average period between price changes of 4.3 months, as estimated by Bils and Klenow (2004). The Taylor rule coefficients, ρ_r , ϕ_π and ϕ_y , are set using estimates in Clarida, Galí, and Gertler (2000).

The final structural parameter, the trade elasticity ϕ , is subject to empirical and theoretical disagreement surrounding its value. Theoretically, Corsetti et al. (2008b) show that both low and high trade elasticities can be consistent with the Backus and Smith (1993) puzzle within a two-country, micro-founded model. Within their model, Corsetti et al. (2008b) find that a low trade elasticity magnifies the consumption risk arising from productivity shocks through terms of trade and real exchange rate volatility. Nevertheless, similar results are obtained with a high trade elasticity when productivity shocks are highly persistent. Empirically, there is a division between macro and micro-estimates of the trade elasticity. Macro-estimates tend to be low; Corsetti et al. (2008b) use the method of moments and estimate the trade elasticity to be around 0.5. Micro-estimates tend to be high; Feenstra, Luck, Obstfeld, and Russ (2014) estimate the trade elasticity to be around 1.5. In the next subsection, I demonstrate that the model most closely matches the stylised facts when the trade elasticity is set to 0.78, closer to the value estimated by Corsetti et al. (2008b).

Shock Process Parameters The parameterisation of the productivity shock a_t mirrors Benigno and Thoenissen (2008), accounting for cross-country correlation. Benigno and Thoenissen (2008) use annual data, so the parameterisation for this model adjusts for this by using quarterly frequency equivalents from Küçük and Sutherland (2015).

The calibration of preference, labour supply, markup, monetary policy and government shock parameters is based on estimates by Smets and Wouters (2005, 2007). The benchmark calibration of the international asset demand shock is due to Itskhoki and Mukhin (2017).

5.4.2 Matching the Stylised Facts

In this sub-section, I discuss the model's ability to match the stylised facts regarding the size and composition of international asset positions summarised in table 5.1. I use the range of average values for the 2002-2008 sub-sample as the target for the model.

I first define the model quantities that I compare to the empirical benchmarks. All quantities are expressed as a fraction of Home country (steady state) GDP, \bar{Y} . Total (net) holdings of Home and US-issued assets in the Home country, t_{HH} and t_{HF} respectively, are defined as:

$$t_{Hk} = \frac{\bar{B}_{Hk} + \bar{B}_{Hk}^L + \bar{S}_{Hk}}{\bar{Y}}, \quad \text{for } k = H, F$$

These summarise the home country's total exposure to all Home and US assets respectively.³⁴ In order to be consistent with the empirical evidence, $t_{HF} \in (0.0554, 0.3366)$.

Home holdings of Home and US-issued equity are given by:

$$s_{HH} = 1 + \frac{\bar{S}_{HH}}{\bar{Y}} \quad \text{and} \quad s_{HF} = \frac{\bar{S}_{HF}}{\bar{Y}}$$

respectively.³⁵ To match the empirical evidence, $s_{HF} \in (0.0270, 0.2071)$. Home holdings of Home and US-issued debt, b_{HH} and b_{HF} respectively, are defined as:

$$b_{Hk} = \frac{\bar{B}_{Hk} + \bar{B}_{Hk}^L}{\bar{Y}}, \quad \text{for } k = H, F$$

which includes both short and long-term bond holdings. To be consistent with the target, $b_{HF} \in (0.0284, 0.1295)$. To further decompose external debt holdings, I define Home holdings of Home and US-issued short and long-term bonds as:

$$b_{Hk}^S = \frac{\bar{B}_{Hk}}{\bar{Y}} \quad \text{and} \quad b_{Hk}^L = \frac{\bar{B}_{Hk}^L}{\bar{Y}}, \quad \text{for } k = H, F$$

respectively. To be consistent with the empirical evidence, $b_{HF}^S \in (0.0065, 0.0192)$ and $b_{HF}^L \in (0.0199, 0.1103)$.

In addition to the stylised facts in section 5.2, there are empirical regularities pertaining to international asset portfolios which the model should also replicate.

First, the equity home bias puzzle describes the fact that external holdings of Foreign equity are lower than predicted by economic theory (French and Poterba, 1991). Agents tend to hold a disproportionately high percentage of domestic equity. To study the model's performance in this respect, I define the home share of Home equity holdings in the Home country equity portfolio \tilde{s}_{HH} as:

$$\tilde{s}_{HH} = \frac{1 + \bar{S}_{HH}}{1 + \bar{S}_{HH} + \bar{S}_{HF}}$$

which has this form because residents initially own 100% of domestic equity before international financial markets open. For the model to account for equity home bias, $\tilde{s}_{HH} \in (0.5, 1)$.

Second, countries have negative external asset positions; they sell domestic equity and bonds

³⁴As the steady state is defined where $\overline{NFA} = 0$, the sum of t_{HH} and t_{HF} is zero for all values of ϕ .

³⁵Home holdings of Home equity have this form because it is assumed that countries initially own 100% of their equity before international financial markets open, as they receive dividends from the home firm in (5.8).

abroad. Within the model, this requires that the Home external position in Home assets is negative: $b_{HH} < 0$, $b_{HH}^S < 0$, and $b_{HH}^L < 0$.

Table 5.3 presents the steady state portfolio quantities for Home investors under the model's baseline calibration from 5.2.³⁶ The PCP model closely matches the stylised facts at the baseline trade elasticity ϕ parameterisation of 0.78. At this value, the model can quantitatively match: total Home holdings of US assets t_{HF} ; equity home bias \tilde{s}_{HH} ; Home holdings of US equity s_{HF} ; Home holdings of Home debt b_{HH} ; Home holdings of Foreign debt b_{HF} ; Home holdings of Home short-term debt b_{HH}^S ; and Home holdings of home long-term debt b_{HH}^L . Although the PCP model does not quantitatively match the Home holdings of US-issued short and long-term bonds, b_{HF}^S and b_{HF}^L respectively, the model is still qualitatively correct; Home households hold positive quantities of US-issued short and long-term bonds in the model's steady state.

Because the trade elasticity ϕ is the parameter in the model with the greatest uncertainty surrounding its value, I also assess the model's ability to match the empirical evidence for different values of ϕ . To do this, I calculate steady state portfolio positions within the model for different values of ϕ ranging from 0.01 to 2.50, in increments of 0.005. This range bounds both the macro (Corsetti et al., 2008b) and micro (Feenstra et al., 2014) estimates presented in the previous sub-section. All other parameters are set to the values defined in table 5.2. Table 5.4 presents the ranges of values of ϕ for which the model can account for each of the stylised facts, implying that it most closely fits the empirical evidence around the baseline calibration.

5.4.3 Determinants of International Asset Positions

In the remainder of this section, I analyse the sources of macroeconomic risk that affect the composition of international asset positions the model, by reporting how steady state asset holdings vary with the volatility of different sources of risk and model parameters. Figure 5.5 plots steady state asset portfolios as a function of the productivity shock volatility σ_a .³⁷ The grey bands depict the range consistent with the empirical evidence regarding international portfolios from table 5.1. The vertical dashed lines are positioned at the baseline value for σ_a .

Although Home agents' steady state holdings of Foreign short-term bonds are positive in the baseline parameterisation, they decrease in the volatility of the productivity shocks, and become negative at sufficiently high values of σ_a . In contrast, Home agents' steady state holdings of Foreign long-term bonds increase with σ_a , but only turn positive when the productivity shock standard deviation is sufficiently high. Additionally, the Home share of Home equity increases with the standard deviation of the productivity shock, while Home holdings of Foreign equity decrease, with both quantities remaining positive for all values of σ_a .

What can explain the relationships between steady state asset holdings and the volatility of productivity shocks? Consider a negative Home productivity shock ($\varepsilon_{a,t} < 0$), which may become increasingly severe as the volatility of productivity shocks increases. Following the shock, relative (habit-adjusted) consumption $C_{X,t}/C_{X,t}^*$ decreases. That is, the marginal utility of Home agents relative to Foreign agents increases. In this state of the world, Home agents would prefer assets that pay out relatively more highly.

This shock has competing effects on the relative returns on Home and Foreign short-term bonds. Because the shock increases the relative price level P_t/P_t^* , as Home goods become

³⁶In appendix D.2, I compare the model's fit under different price-setting regimes.

³⁷Shock volatilities are kept the same in the two countries at all times.

Table 5.3: Steady State Portfolio Quantities for Home Investors Under the Model's Baseline Calibration

HOLDINGS	DATA	PCP
Total US Assets, t_{HF}	5.54-33.66%	5.68%
Equity Home Bias, \tilde{s}_{HH}	50-100%	97.09%
US Equity, s_{HF}	2.70-20.71%	2.77%
Home Debt, b_{HH}	<0%	− 3.71%
US Debt, b_{HF}	2.84-12.95%	2.90%
Home Short-Term Debt, b_{HH}^S	<0%	− 2.79%
US Short-Term Debt, b_{HF}^S	0.65-1.92%	<i>1.97%</i>
Home Long-Term Debt, b_{HH}^L	<0%	− 0.93%
US Long-Term Debt, b_{HF}^L	1.99-11.03%	<i>0.93%</i>

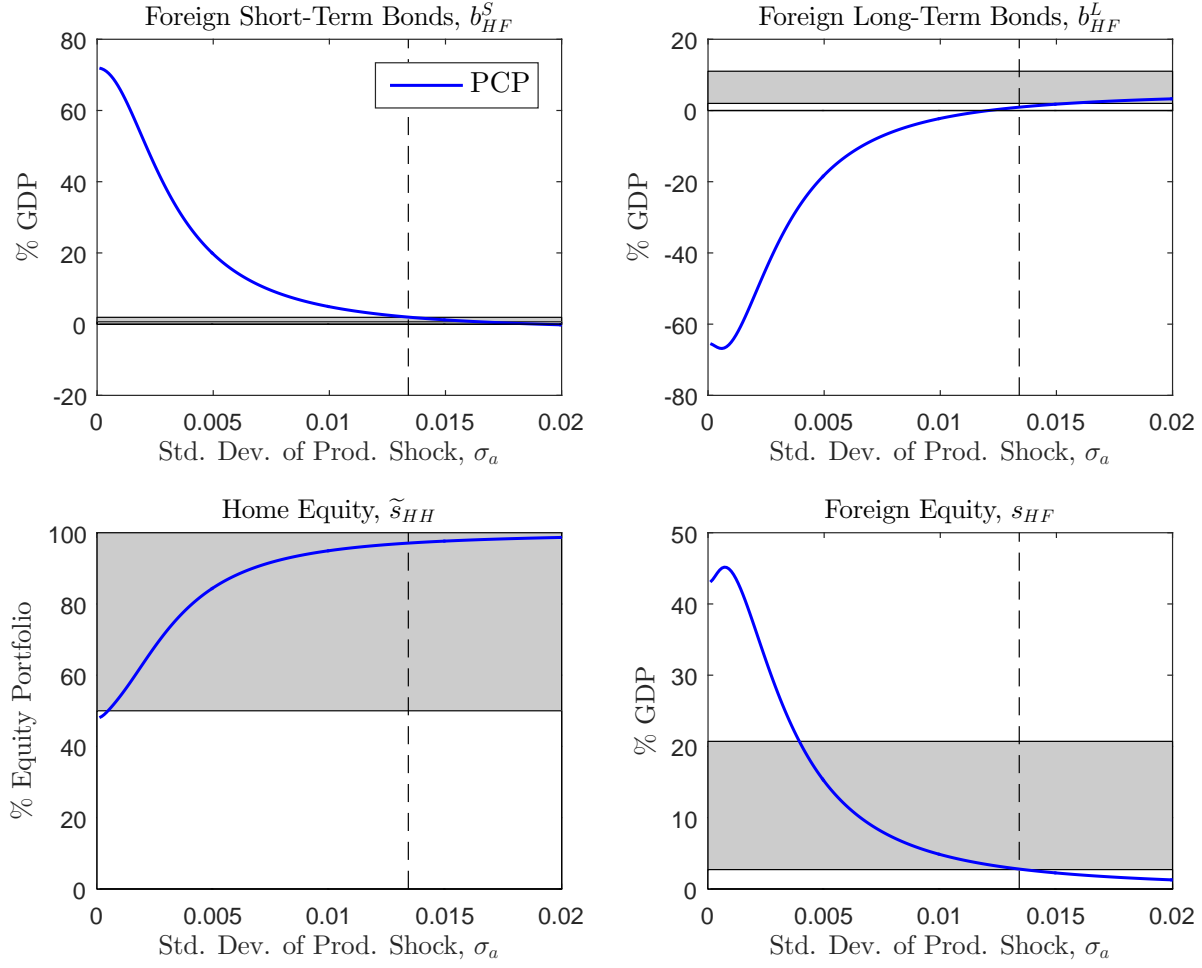
Notes: Steady state portfolio quantities under the baseline model calibration with $\phi = 0.78$. Emboldened values denote quantities that quantitatively match the empirical evidence. Italicised values denote quantities that qualitatively match the empirical evidence. Values that are neither emboldened nor italicised do match the empirical evidence qualitatively or quantitatively.

Table 5.4: Ranges of Values of the Trade Elasticity ϕ within which Steady State Portfolio Quantities for Home Investors are Consistent with Empirical Evidence

HOLDINGS	PCP
Total US Assets, t_{HF}	0.690-0.785
Equity Home Bias, \tilde{s}_{HH}	0.185-0.205, 0.640-0.790
US Equity, s_{HF}	0.725-0.785
Home Debt, b_{HH}	0.005-0.880
US Debt, b_{HF}	0.325-0.380, 0.775-1.295
Home Short-Term Debt, b_{HH}^S	0.005-0.790
US Short-Term Debt, b_{HF}^S	0.780-0.785
Home Long-Term Debt, b_{HH}^L	0.775-2.500
US Long-Term Debt, b_{HF}^L	0.780-0.805

Notes: Ranges of values of the trade elasticity ϕ within which the steady state portfolio quantities of the model, presented in section 5.3, are consistent with the empirical evidence presented in section 5.2.2. I solve the model for values of ϕ ranging from 0.005 to 2.500, in increments of 0.005.

Figure 5.5: International Asset Portfolios in the PCP Model's Steady State and the Volatility of Productivity Shocks σ_a



Note: Home holdings of Foreign assets ($b_{HF}^S, b_{HF}^L, s_{HF}$), and Home share of Home equity \tilde{s}_{HH} , in the steady state of the PCP model plotted against the volatility of productivity shocks σ_a . The volatility is the same for home and foreign economies, and is plotted at increments of 0.0001 from $\sigma_a = 0.0001$ to $\sigma_a = 0.02$. The values of all other parameters are defined in table 5.2. The grey shaded area denotes the area consistent with the empirical evidence presented in section 5.2.2. The vertical dashed line is positioned at $\sigma_a = 0.0134$.

relatively more scarce, it erodes the relative real return on Home bonds. However, the real exchange rate Q_t will also immediately appreciate, reducing the return on Foreign short-term bonds in units of the Home consumption basket. For Home agents' steady state holdings of short-term bonds to be positive in the baseline calibration, the former of these effects must dominate. However, the latter effect explains why Home holdings of Foreign short-term bonds turn negative at sufficiently high values for σ_a . The immediate exchange rate appreciation that a negative Home productivity shock induces will, *ceteris paribus*, increase the relative return on Home short-term bonds at precisely the time when relative Home (habit-adjusted) consumption is low, such that Foreign short-term bonds will not provide a good hedge for productivity shocks.

When σ_a is sufficiently high in comparison to other macroeconomic shocks, Foreign long-term bonds provide a good hedge for productivity shocks. The same negative Home productivity shock will generate a persistent increase in relative Home inflation $\hat{\pi}_t - \hat{\pi}_t^*$, which is expected to erode the value of Home long-term bonds relative to Foreign long-term bonds over their lifetime. Consequently, the value of Home long-term bonds $P_{L,t}$ will fall after a negative Home productivity shock and Home investors will view Foreign long-term bonds as a better hedge

for productivity shocks *ex ante*, because their relative resale price is high at precisely the time when relative (habit-adjusted) consumption is low. Therefore, steady state holdings of Foreign long-term bonds by Home households are increasing in the volatility of productivity shocks.

The relationship between equity holdings and the volatility of the productivity shock can be explained by the reaction of relative non-financial income $w_t L_t / w_t^* L_t^*$ to productivity shocks, as in Coeurdacier and Gourinchas (2016). The mechanism in this model relies on price stickiness, mirroring Engel and Matsumoto (2009). Consider a positive Home productivity shock. Firms who are able to reset their price in the period of the shock will cut it, reducing relative Home inflation $\hat{\pi}_t - \hat{\pi}_t^*$. However, Home firms who are unable to reset their price will be forced to economise on labour to reduce costs. For this reason, relative labour earnings in the Home economy $w_t L_t / w_t^* L_t^*$ will fall. To hedge this change in their non-financial income, investors will seek assets that pay relatively highly in this state of nature. Home equity provides a hedge for this, conditional on bond returns hedging the change in relative (habit-adjusted) consumption, because the reduction in Home labour costs will increase the relative profitability of home firms. Engel and Matsumoto (2009) discuss how this mechanism can generate equity home bias within their model; the same reasoning can be applied here.

Importantly, because the model is symmetric, the above reasoning is invariant to the consideration of Foreign shocks. For instance, a positive Foreign productivity shock will reduce relative Home (habit-adjusted) consumption $C_{X,t} / C_{X,t}^*$ and increase relative Home labour income $w_t L_t / w_t^* L_t^*$. Home households will seek bonds that pay out highly in this state of nature to hedge changes in relative (habit-adjusted) consumption, but will prefer equity that pays out relatively more when relative labour income is low. Foreign long-term bonds continue to provide a good hedge for the former risk, while Home short-term bonds will not. The positive foreign productivity shock will reduce prices abroad, increasing P_t / P_t^* . The immediate Home real exchange rate Q_t appreciation will reduce the return on Foreign short-term bonds in units of the Home consumption basket, while the persistent increase in relative Home inflation $\hat{\pi}_t - \hat{\pi}_t^*$ will erode the relative value of Home long-term bonds. Because of this, investors will use Foreign long-term bonds to insure against this shock *ex ante*, while holding fewer Foreign short-term bonds in the model's steady state. Home equity continues to provide a good hedge for fluctuations in non-financial income, because relative labour income and relative firm profitability are inversely correlated.

5.5 Global Transmission Through Long-Term Interest Rates

In this section, I use the model from section 5.3 to study the global spillover effects of US monetary policy, building on existing literature by isolating a specific transmission channel through longer-term interest rates and international bond markets.

5.5.1 Model Predictions

I use the model to investigate the impact responses of interest rates to a US (i.e. Foreign) monetary policy shock $\varepsilon_{mp,t}^*$. Although the bond duration in the baseline calibration is five years, I study the response of the whole term structure of interest rates. Within the model, this is possible because long-term bonds are defined as perpetuities with exponentially declining coupon payments κ . I present impact responses for values of κ set so that the long-term bond

duration ranges from one year ($L = 4$ quarters) to ten years ($L = 40$ quarters), with all other parameters maintained at their baseline values.

Figure 5.6 plots the model-implied impact responses of interest rates, asset positions and the exchange rate to a surprise US monetary policy tightening of approximately 100 basis points. The horizontal axis denotes the long-term bond duration in years. It demonstrates that a surprise increase in the US monetary policy rate unambiguously increases US long-term interest rates at all durations. It also illustrates that the US policy rate change has global effects. First, in accordance with conventional wisdom, the Home currency depreciates on impact, in nominal and real terms. This is independent of the long-term bond duration. Second, Home long-term interest rates respond to the US monetary policy shock, increasing for each duration; the global transmission of US monetary policy through longer-term interest rates serves to amplify comovements between advanced economies.

What generates these spillover effects via longer-term interest rates? Following a tightening of US monetary policy, the Home exchange rate will depreciate. This will generate imported inflation in the Home country, as the Home currency price of US imports increases, necessitating a tightening of Home monetary policy in the future. In turn, this raises expectations of future Home short-term interest rates, placing upward pressure on Home long-term interest rates. However, because Home monetary policy is expected to tighten in the future, the Home real exchange rate, which depreciates immediately after the US monetary tightening, will be expected to appreciate in the future as the world economy returns to its steady state. Because Home investors expect this real exchange rate appreciation, which will reduce the relative return on US long-term bonds in Home consumption units, they will rebalance their portfolio towards Home long-term bonds. That is, the Home advanced economy does not suffer from a capital outflow through the bond market.³⁸ This portfolio rebalancing will bid down Home longer-term interest rates, serving to offset the increase in Home longer-term interest rates due to increased interest rate expectations to some extent. Bond portfolio rebalancing in advanced economies serves to attenuate some of the spillover effects of US monetary policy that transmit through longer-term interest rates.

5.5.2 Empirical Comparison

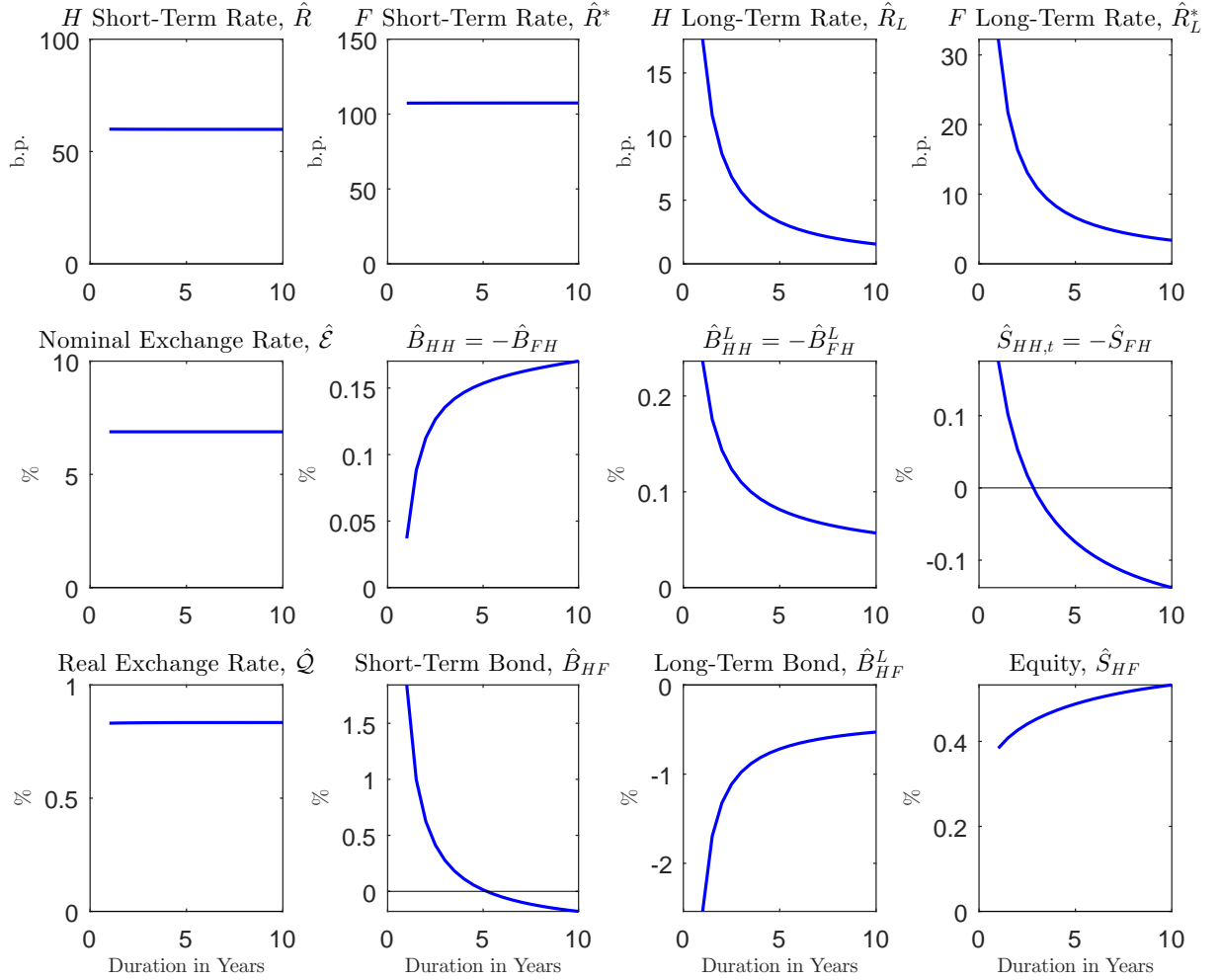
I compare the model's predictions to empirical evidence by carrying out an event study into the impact responses of longer-term interest rates in Australia, Canada, France, Germany and the UK to US monetary policy surprises. I use daily frequency interest rate data, and isolate US monetary policy surprises using intraday data, to estimate the following regression:

$$\Delta y_{L,t-1,t}^{(k)} = \alpha_L + \beta_L mp_t + u_{L,t} \quad (5.43)$$

where $\Delta y_{L,t-1,t}^{(k)}$ is the change in the country- k L -quarter (net) interest rate from day $t - 1$ to t , mp_t denotes the surprise change in US monetary policy on day t , and $u_{L,t}$ is a disturbance term. α_L represents the average daily change in the L -quarter interest rate on days without monetary policy surprises, and β_L represents the impact response of the interest rate to a surprise 100 basis point tightening of US monetary policy, equivalent to the impact responses in figure 5.6.

³⁸Dedola et al. (2016) reach a similar conclusion for advanced economies. Within an empirical vector autoregression study, these authors find that advanced economies do not see capital outflow following a US monetary policy tightening, although emerging market economies do.

Figure 5.6: PCP Model-Implied Impact Response of Interest Rates, the Exchange Rate and International Portfolios to a Surprise US (Foreign F) Monetary Policy Tightening of Approximately 100 Basis Points



Note: Impact response of Home (Foreign) short-term interest rates $\hat{R}_t^{(*)}$, long-term interest rates $\hat{R}_{L,t}^{(*)}$, international asset holdings, and the nominal and real exchange rate ($\hat{\mathcal{E}}_t$ and \hat{Q}_t) to a surprise tightening, of approximately 100 basis points (b.p.), in US monetary policy $\varepsilon_{mp,t}^*$. $\hat{B}_{Hk,t}$ ($\hat{B}_{Hk,t}^L$) denotes the impact response of home holdings of home (when $k = H$) and foreign (when $k = F$) short-term (long-term) bonds. $\hat{S}_{HH,t}$ ($\hat{S}_{HF,t}$) denotes the impact response of home holdings of home (foreign) equity. The impact responses are plotted for variants of the model with different long-term bond durations, ranging from $L = 4$ (one year) to $L = 40$ (ten years) where κ is altered to match the required long-term bond duration. The values of all other parameters are defined in table 5.2. ‘PCP’ denotes producer currency pricing. All interest rate responses are annualised.

Data

I estimate (5.43) using daily frequency zero-coupon government bond yields as the dependent variable. The duration and maturity of zero-coupon bonds are identical, permitting comparison with the model impact responses. Specifically, I use the following maturity yields: 1 year, 18 months, 2 years, 30 months, 3 years, 42 months, 4 years, 54 months, 5, 7 and 10 years. The sample of countries is the same as in section 5.2, with the US acting as the base country from where the monetary policy shock emanates.³⁹

I measure US monetary policy surprises using intraday movements in the current calendar month federal funds futures (FFFs) rates in a thirty-minute window around Federal Open Market Committee (FOMC) announcements, compiled by Gürkaynak et al. (2005a). Changes in the current month FFFs rate, adjusted for the timing of the announcement within the month, can be associated with revisions in expectations of the effective federal funds rate for the remainder of the month and measure the surprise component of the FOMC decision (Kuttner, 2001).

I estimate (5.43) using data from January 2002 to December 2015 for all six countries (118 announcement days). I choose this start date for reasons stated in section 5.5.3. The sample ends in December 2015 because the monetary policy surprise series concludes in this month.

Because the daily zero-coupon bond yields are quoted at the closing time of relevant markets, I adjust the data for the event study. For instance, by the time an FOMC announcement occurs at 2.15pm in the US, the Australian, French, German and UK markets will have closed. Therefore, the relevant daily change in yields in these jurisdictions comes on the calendar date after the US announcement.

Event Study

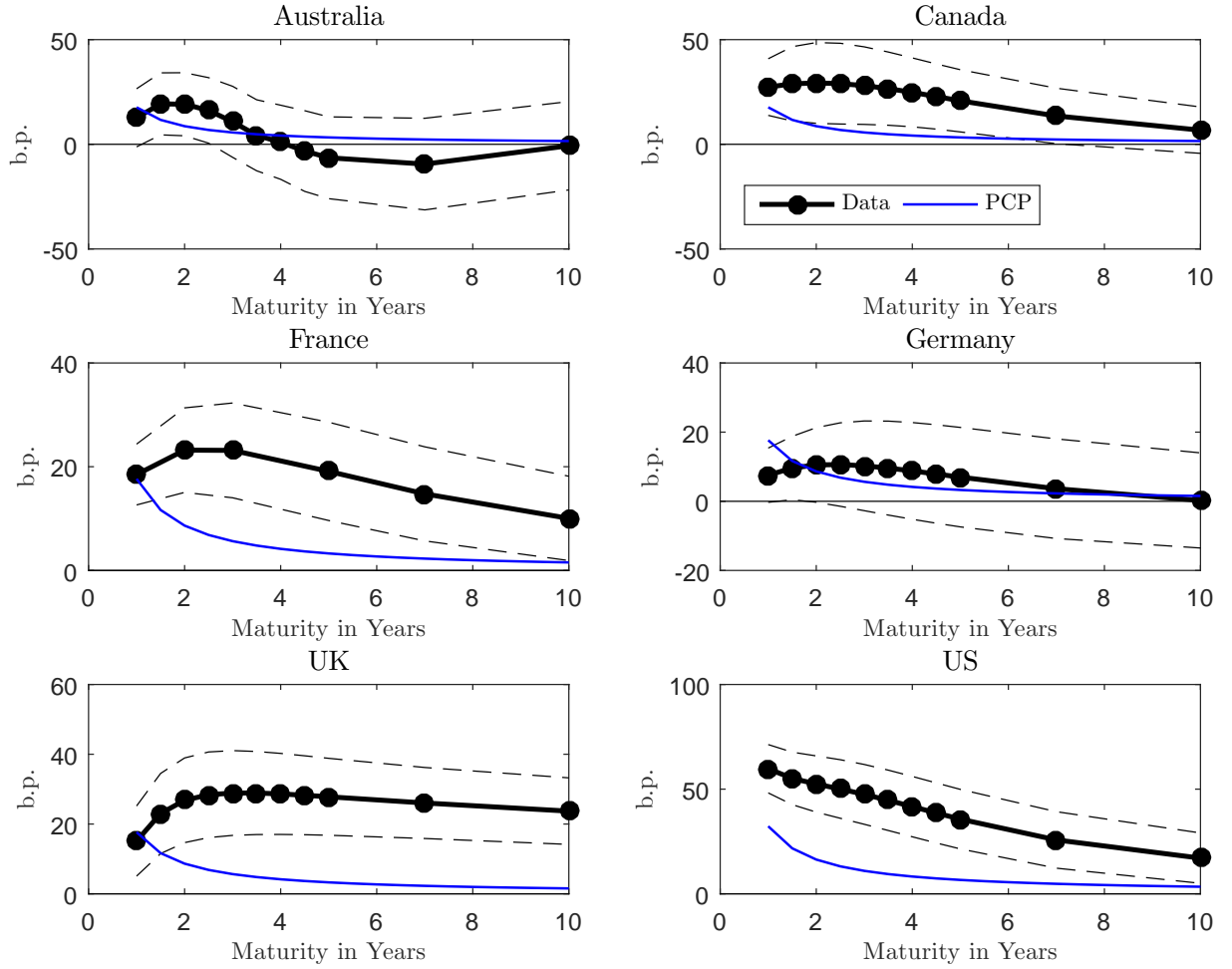
Figure 5.7 presents the impact response of bond yields to a surprise 100 basis point US monetary policy tightening. The black dots represent the estimated $\hat{\beta}_L$ coefficients and the thin dashed lines correspond to 95% confidence intervals using heteroskedasticity and autocorrelation robust Newey and West (1987) standard errors. For comparison, the thin blue line denotes the impact responses of equivalent interest rates to an equal-sized shock plotted in figure 5.6.

The bottom-right panel of figure 5.7 illustrates that, following a 100 basis point surprise tightening of US monetary policy, US longer-term interest rates significantly increase at all maturities. The magnitude of the increase is monotonically decreasing with the maturity, with the 1-year yield responding most strongly.

The remaining panels of figure 5.7 indicate that many longer-term interest rates in other advanced economies increase around a surprise US monetary policy tightening. In France and the UK, interest rates significantly increase at all horizons on FOMC announcement days. In the UK, the 3-year yield is most responsive, increasing by 28.9 basis points around a 100 basis point surprise US tightening. Canadian bond yields are also sensitive to US monetary policy surprises, significantly increasing at all horizons out to 7 years. The 2-year Canadian yield is the most responsive of all plotted in figure 5.7, increasing by 29.3 basis points around a 100 basis point surprise US tightening. Point estimates of the responsiveness of German bond yields are positive at all horizons, albeit only statistically significant out to 2 years. Interestingly, the response of German bond yields is about half that of French yields. Australian bond yields react

³⁹Because of data availability, I use fewer maturities for the French results: 1, 2, 3, 5, 7 and 10 years. A complete description of data sources is in appendix A.1.

Figure 5.7: Impact Response of Bond Yields to a Surprise US Monetary Policy Tightening of 100 Basis Points Compared to PCP Model-Implied Impact Response



Note: The black dots denote the impact response of (annualised) longer-term interest rates of different maturities in Australia, Canada, France, Germany, the UK and the US to a surprise US monetary tightening of 100 basis points, corresponding to the estimated β_L coefficients in equation (5.43). The horizontal axis denotes the maturity in years. Interest rates are zero-coupon government bond yields. The monetary policy surprise is measured using the intraday movement in the current month federal funds futures rate in a thirty minute window around an FOMC announcement (Gürkaynak et al., 2005a). The dashed lines represent the 95% confidence intervals around the estimated coefficients, constructed using robust standard errors. The estimates are constructed using data from January 2002 to December 2015. The thin blue line denotes the corresponding impact responses of the equivalent interest rates to an equal-sized shock from the PCP model laid out in section 5.3.

most uniquely to a surprise tightening of US monetary policy. The estimated coefficients are positive for the 1 to 4-year maturities, but turn negative — albeit statistically insignificant — at longer maturities.

Visually, the response of the German yield curve most closely matches the predictions of the PCP model. The model-implied quantities lie within the estimated confidence intervals at all maturities beyond 1-year. In Canada, the model-implied predictions lie within the estimated confidence intervals at maturities below 2 years and in excess of 5 years, while in France and the UK, the model-implied quantities lie below the estimated confidence intervals at maturities of 2 years or more. Taken together, the model-implied responses are qualitatively similar to those implied by the data in Canada, France, Germany and the UK, but indicate that the international spillover effects of US monetary policy through longer-term interest rates may be stronger than implied by the model.

5.5.3 Long-Term Interest Rate Decomposition

To understand why the global macroeconomic spillover effects of US monetary policy through longer-term interest rates is stronger than implied by the model, I use the canonical decomposition of longer-term interest rates into a risk-neutral expected future short-term interest rate component and a term premium (4.1), where $y_{L,t}$ is the yield on an L -period government bond at time t ,⁴⁰ $y_{1,t}$ is the one-period (net) interest rate, and $tp_{L,t}$ is the term premium on the L -period rate. The first term on the right-hand side of (4.1), $exp_{L,t} \equiv \frac{1}{L} \mathbb{E}_t \sum_{l=0}^{L-1} y_{1,t+l}$, defines the average of expected future short-term interest rates between period t and the bond's maturity.

This decomposition lends itself to the study of international spillovers of monetary policy for a number of reasons. First, the decomposition is of direct relevance to policy as the two sub-components have differing policy implications. If, following a US monetary policy announcement, long-term interest rate movements in other advanced economies are associated with changes in interest rate expectations, then monetary policymakers in other advanced economies may attenuate these spillover effects by clearly communicating to investors through policies such as forward guidance. Insofar as investors view different asset classes and maturities as imperfect substitutes, the price of an asset will rise and its term premium fall when demand for that asset increases. If US policy influences longer-term interest rates in other advanced economies through term premia, then this motivates a need for policymakers to keep a watchful eye on international capital flows and the pricing of risk.

Second, the decomposition in (4.1) can be estimated empirically, using a no-arbitrage Gaussian affine dynamic term structure model (GADTSM).

I use daily frequency estimates of expectations and term premia to extend the event study from the previous subsection by estimating the following regressions:

$$\Delta exp_{L,t-1,t}^{(k)} = \alpha_{L,e} + \beta_{L,e} mp_t + u_{L,e,t} \quad (5.44)$$

$$\Delta tp_{L,t-1,t}^{(k)} = \alpha_{L,tp} + \beta_{L,tp} mp_t + u_{L,tp,t} \quad (5.45)$$

where $exp_{L,t-1,t}^{(k)}$ is the change in the average expectation of the short-term interest rate over the next L quarters in country k from day $t-1$ to t , and $\Delta tp_{L,t-1,t}^{(k)}$ is the daily change in the L -quarter term premium. These regressions extend upon the results in section 5.5.2 by focusing on

⁴⁰In comparison to the gross long-term interest rate defined within the theoretical model, $R_{L,t} \equiv 1 + y_{L,t}$.

policy-relevant mechanisms through which US monetary policy can exert international spillover effects through longer-term interest rates.

I estimate interest rate expectations and term premia by applying the OIS-augmented no-arbitrage GADTSM proposed in chapter 3. I use the same daily frequency government bond yield data as in section 5.5.2. For GADTSM-estimation, I also use 3 and 6-month interest rates.⁴¹ In addition, I use daily frequency 3, 6, 12 and 24-month OIS rates, where available. The choice of maturities is motivated by evidence in chapter 2 suggesting that OIS rates with maturities out to 2 years accurately reflect investors' expectations of future short-term interest rates. However, because of data availability, I only use the 3, 6 and 12-month OIS rates for the Australian and Canadian decompositions. The sample runs from January 2, 2002 to December 31, 2015, starting in 2002 because OIS rate data is not available from 1999 in most countries, with many series beginning in 2001 or 2002. This sample period is the same as in chapter 4 for the US.⁴² I plot the OIS-augmented GADTSM for the US in section 4.3, and demonstrate that it provides a superior fit of corresponding-horizon FFFs and survey expectations.⁴³

Event Study Results

Endowed with daily frequency estimates of interest rate expectations and term premia for each of the six countries, I estimate (5.44) and (5.45). I plot the impact responses of bond yield components for maturities of 1 year or more.⁴⁴

Figure 5.8 presents estimates of the responsiveness of interest rate expectations to a surprise 100 basis point tightening of US monetary policy from (5.44). The bottom-right panel of figure 5.8 illustrates that this significantly increases US interest rate expectations at all horizons.

The remaining panels illustrate the response of the term structure of interest rate expectations in the other five advanced economies. In Canada, France, and the UK, a surprise tightening of US monetary policy is associated with an immediate and significant, increase in interest rate expectations at all horizons, consistent with investors forming expectations that other advanced economies will tighten monetary policy following a US monetary policy tightening. Moreover, in these countries, the impact response of the term structure of interest rate expectations is hump-shaped. Interest rate expectations are most responsive in these economies at the 3.5 to 5-year horizons. This suggests that international spillover effects of US monetary policy through interest rate expectations are pertinent at a wide range of horizons, extending beyond the near-term. In Australia and Germany, the estimated responses of interest rate expectations are also positive at all horizons, albeit statistically insignificant at most tenors.⁴⁵ At long horizons espe-

⁴¹See appendix A.1 for further information on data sources.

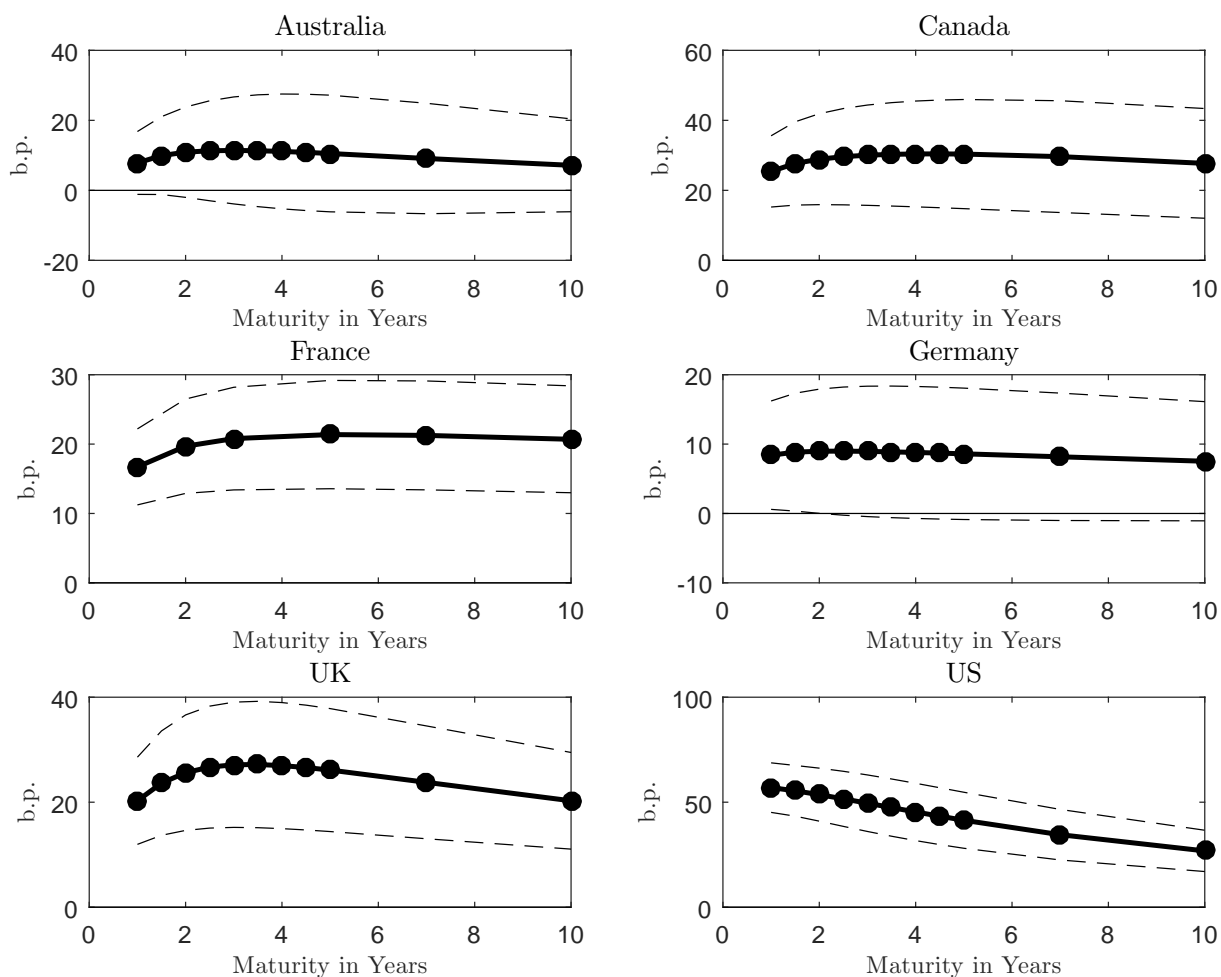
⁴²A direct comparison between the GADTSM and the theoretical model cannot be made because the asset pricing kernel within the theoretical model is not consistent with the GADTSM presented in chapter 3.

⁴³Estimation results for the remaining five countries are discussed in appendix D.3.

⁴⁴I also carry out two, unreported, robustness exercises. In the first, I account for the possibility that US 'unconventional' monetary policies, enacted since late-2008, may influence the conclusions made here about conventional monetary policy. I find that the transmission of US monetary policy to other advanced economies through the term structure of interest rates in 2002-2008 was not significantly different from 2002-2015. In the second, I find that the results are not significantly different when an unaugmented GADTSM is used.

⁴⁵In figure 5.8, the response of German interest rate expectations is about half that of French interest rate expectations. This occurs because German and French yield curve decompositions are estimated independently. An alternative approach to estimating Eurozone decompositions, which accounts for commonality of monetary policy rates, involves using estimating interest rate expectations with bond yields from one country (i.e. Germany) and calculating term premia for the other using (4.1). If this method is applied, the impact response of French interest rate expectations would be the same as German expectations in figure 5.8, while the impact response of French term premia would be higher at all maturities than in figure 5.9.

Figure 5.8: Impact Response of Interest Rate Expectations to a Surprise US Monetary Policy Tightening of 100 Basis Points



Note: The black line denotes the impact response of interest rate expectations of different maturities in Australia, Canada, France, Germany, the UK and the US to a surprise US monetary tightening of 100 basis points, corresponding to the estimated $\beta_{L,e}$ coefficients in equation (5.44). The horizontal axis denotes the maturity of the interest rate expectations in years. The interest rate expectations are estimated using the OIS-augmented GADTSM of Lloyd (2017a). The monetary policy surprise is measured using the intraday movement in the current month federal funds futures rates in a thirty minute window around an FOMC announcement (Gürkaynak et al., 2005a). The dashed lines represent the 95% confidence intervals around the estimated coefficients, constructed using robust standard errors. The estimates are constructed using data from January 2002 to December 2015.

cially, the impact responses of interest rate expectations tend to be greater than the responses of corresponding-maturity bond yields, indicating that spillovers through this channel can explain the differences between empirical and model-implied responses in figure 5.7.

Figure 5.9 presents the impact responses of term premia to a surprise tightening of US monetary policy. The bottom-right panel of the figure depicts the response of US term premia to a US monetary policy tightening. At all horizons, US term premia fall significantly following a surprise tightening of US monetary policy. The largest fall in US term premia occurs at the 4-year horizon; following a surprise 100 basis point tightening of US monetary policy the premium at this tenor declines by 7.57 basis points.

The remaining panels of figure 5.9 plot the responsiveness of the term structure of term premia in Australia, Canada, France, Germany and the UK to a surprise US monetary policy tightening. In Australia, Canada, France and Germany, the longer-horizon term premia fall significantly in response to a surprise tightening of US monetary policy.⁴⁶ Moreover, the responses tend to increase in magnitude at longer horizons, suggesting that the international spillover effects of US monetary policy through term premia are most pertinent at longer-term horizons.

Taken together, the results in figures 5.8 and 5.9 indicate that US monetary policy announcements do have spillover effects to other advanced economies that propagate through longer-term interest rates. Moreover, the decomposition of longer-term interest rates in (4.1) sheds further light on these spillovers. Two implications are noteworthy. First, movements in interest rates in many advanced economies on FOMC announcement dates are strongly positively associated with changes in interest rate expectations at horizons beyond the near-term. This suggests the monetary policymakers in advanced economies may be able to attenuate some of the spillover effects of US monetary policy by clearly communicating expectations of future short-term interest rates through policies such as forward guidance at a range of horizons. Second, at longer horizons, FOMC announcements are associated with reductions in term premia in other advanced economies that serve to attenuate some of the spillover effects of US monetary policy, and are consistent with the portfolio rebalancing of investors seen in the model in figure 5.6, where investors rebalance their portfolio towards Home long-term bonds because they expect future real exchange rate appreciation.

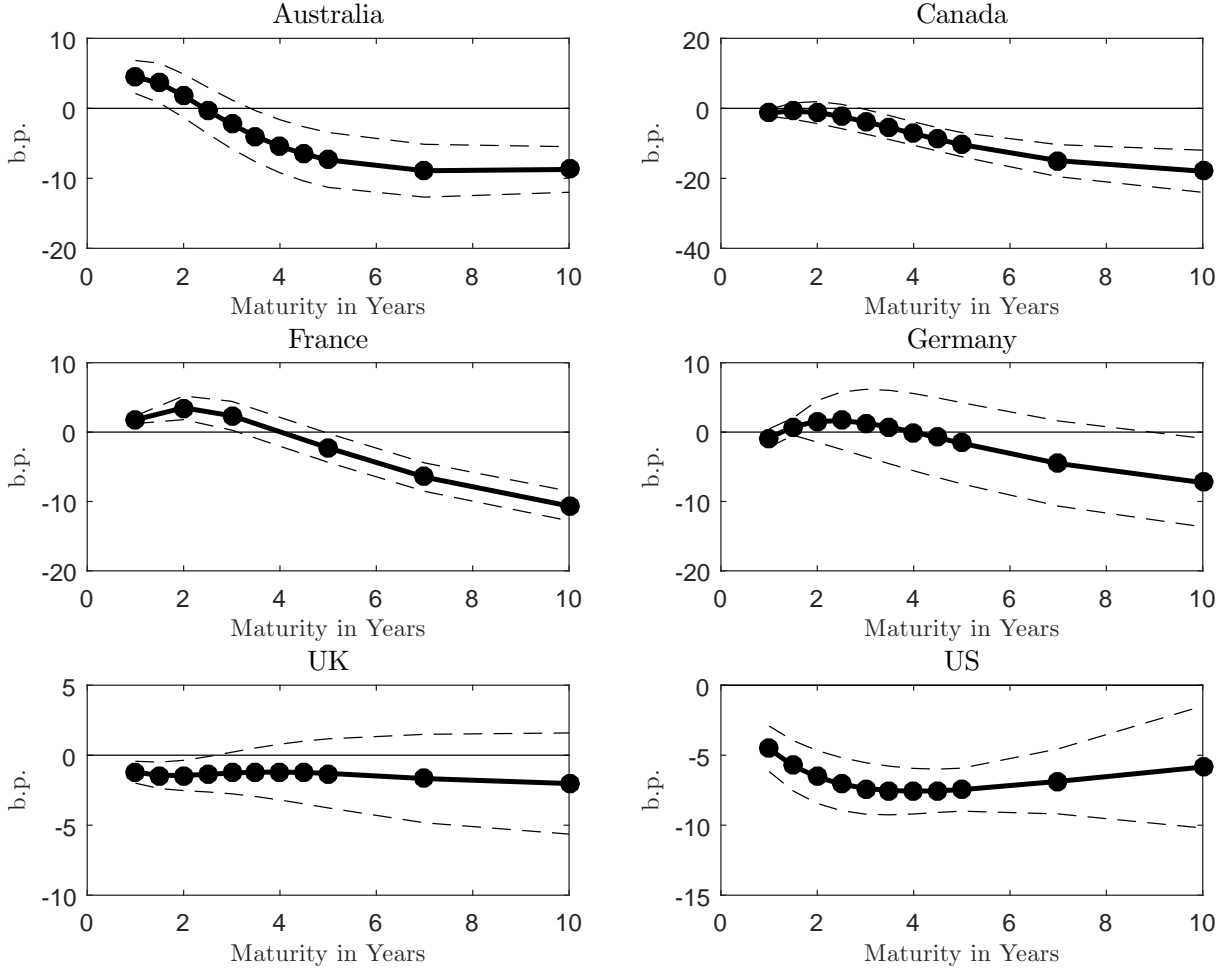
This conclusion echoes a finding of Stavrakeva and Tang (2016) who assess quarterly frequency correlations between longer-term interest rates and exchange rates amongst advanced economies. They find that correlations between shorter-term interest rates and exchange rates are predominantly driven by interest rate expectations, and correlations between longer-term interest rates and exchange rates by term premia. Insofar as US monetary policy surprises influence exchange rates, the empirical results in this chapter suggest that US monetary policy shocks may be one of the drivers of the correlations found in Stavrakeva and Tang (2016).

5.6 Conclusion

In this chapter, I study the mechanisms through which international macroeconomic spillovers between advanced economies propagate through longer-term interest rates. I present a micro-founded, two-country model with endogenous portfolio choice amongst country-specific equity,

⁴⁶In the UK, the point estimates for the response of term premia are negative at all horizons, albeit statistically insignificant from the 3-year maturity onwards.

Figure 5.9: Impact Response of Term Premia to a Surprise US Monetary Policy Tightening of 100 Basis Points



Note: The black line denotes the impact response of term premia at different maturities in Australia, Canada, France, Germany, the UK and the US to a surprise US monetary tightening of 100 basis points, corresponding to the estimated $\beta_{L,tp}$ coefficients in equation (5.45). The horizontal axis denotes the maturity of the term premia in years. The term premia are estimated using the OIS-augmented GADTSM of Lloyd (2017a). The monetary policy surprise is measured using the intraday movement in the current month federal funds futures rates in a thirty minute window around an FOMC announcement (Gürkaynak et al., 2005a). The dashed lines represent the 95% confidence intervals around the estimated coefficients, constructed using robust standard errors. The estimates are constructed using data from January 2002 to December 2015.

short and long-term bonds. The model provides novel insights about the different roles played by short and long-term bonds in international risk sharing and the global transmission of monetary policy shocks. Within the model, short-term bonds are predominantly used to hedge real exchange rate fluctuations that occur immediately after a macroeconomic shock, while long-term bonds mainly hedge expected inflation and real exchange rate movements.

I assess the model's predictions regarding the international transmission of US conventional monetary policy shocks through longer-term interest rates. Within the model, a surprise tightening of US monetary policy generates immediate increases in longer-term interest rates in the US and other advanced economies, that qualitatively align with estimates from an event study, indicating that US monetary policy has powerful global spillover effects through longer-term interest rates.

I extend the empirical analysis by estimating a decomposition of longer-term interest rates into interest rate expectations and term premia. I find that US monetary policy shocks exert powerful spillover effects through interest rate expectations. Surprise US monetary policy tightening tends to immediately increase investors' expectations of future short-term interest rates in other advanced economies at a range of horizons beyond the near-term. In contrast, following the same surprise, term premia in other advanced economies tend to fall, especially at longer horizons, attenuating the global spillover effects of monetary policy.

These findings have important implications for monetary policymakers in advanced economies and contribute to the growing literature on the global financial cycle (Rey, 2014; Miranda-Agrippino and Rey, 2015; Passari and Rey, 2015; Rey, 2016). Monetary policymakers may be able to contain spillover effects through interest rate expectations by clearly communicating their intentions for future policy through policies such as forward guidance at a range of horizons. Nevertheless, the reaction of term premia help to attenuate the spillover effects of US monetary policy to other advanced economies. These movements warrant careful consideration of international capital flows and the pricing of risk, as they indicate that global portfolio rebalancing can help to circumvent the Mundellian trilemma.

Appendix A

Data Sources

A.1 Data Sources

The below table lists sources for US financial market data.

Data Series	Source
US Effective Federal Funds Rate and Treasury Bill Rates	Federal Reserve Statistical Release H.15. Used in chapters 2-5.
US Zero-Coupon Treasury Yields	Gürkaynak, Sack, and Wright (2007a), Federal Reserve. Used in chapters 2-5.
US OIS Rates	<i>Bloomberg</i> . Used in chapters 2-5.
US Federal Funds Futures	<i>Bloomberg</i> and <i>Quandl</i> . Used in chapters 2-3.
US Survey Forecasts of the 3-Month T-Bill Rate	<i>Survey of Professional Forecasters</i> , Federal Reserve Bank of Philadelphia. Used in chapters 2-5.
US Merrill Lynch Option Volatility Estimate (MOVE)	<i>Bloomberg</i> , with the code MOVE Index . This is a yield curve weighted index of the normalised implied volatility on one month Treasury options. It is a weighted average of volatilities on the current US 2, 5, 10 and 30 year government notes. Used in chapter 3.
US Survey-Augmented GADTSM Decomposition	Kim and Wright (2005), Federal Reserve. Used in chapter 3.

The below table lists sources for monthly frequency US macroeconomic data used in chapter 4.

Data Series	Source
US Industrial Production	St. Louis FRED.
US Consumer Price Index	US Bureau of Labor Statistics (BLS).
US Bank Credit	Federal Reserve, table H.8, <i>Assets and Liabilities of Commercial Banks in the US</i> .
US Real Exchange Rate	Broad effective real exchange rate for the US against 60 other economies from the Bank of International Settlements (BIS).

The subsequent table presents sources of financial market data for countries other than the US.

Data Series	Source
Australian Zero-Coupon Government Bond Yields	Reserve Bank of Australia. Used in chapter 5.
Australian OIS Rates	<i>Bloomberg</i> . Used in chapter 5.
Canadian Zero-Coupon Government Bond Yields	Bank of Canada. Used in chapter 5.
Canadian OIS Rates	<i>Bloomberg</i> . Used in chapter 5.
Eurozone OIS Rates	<i>Bloomberg</i> . Used in chapters 2 and 5.
Eurozone EONIA	European Central Bank. Used in chapter 2.
French Treasury Bill Rates and Zero-Coupon Government Bond Yields	<i>Bloomberg</i> and <i>Datastream</i> . Used in chapter 5.
German Zero-Coupon Government Bond Yields	Bundesbank. Used in chapter 5.
Japanese OIS Rates	<i>Bloomberg</i> . Used in chapter 2.
Japanese TONAR	Bank of Japan. Used in chapter 2.
UK Zero-Coupon Government Bond Yields	Bank of England. Used in chapters 2 and 5.
UK OIS Rates	<i>Bloomberg</i> . Used in chapters 2 and 5.
UK SONIA	Bank of England. Used in chapter 2.

A.1.1 Availability of OIS Rate Data

In each jurisdiction, the availability of daily OIS rate data varies with the maturity of the contract. I document the data availability below.

Australia For the 1, 2, 3, 4, 5, 6, 9 and 12-month maturities, the data series start in January 2002; 7, 8, 10 and 11-month OIS rate data is available from May 2002; 18 and 24-month OIS rate data is available from June 2003; and 36-month data begins in April 2004.

Canada For the 1, 2, 3, 4, 5, 6 and 9-month maturities, the data series start in May 2002. Although 1-year OIS rates are available June to October 2003, they are not available for all business days until October 2007. 7, 8, 10, 11, 18 and 24-month OIS rates are only available from October 2007, while 3, 4 and 5-year OIS rates are first available from January 2011.

Eurozone Remolana and Wooldridge (2003) document the growth of the Eurozone OIS market since the inception of the Euro in 1999 to 2003. 1, 2, 3, 4, 5, 6, 7, 8, 9, 10, 11 and 12-month OIS rates are available from January 3, 2000, at least. 15 and 21-month OIS rates are available at a daily frequency from August 22, 2001. Although observations for 18 and 24-month OIS rates are available from as early as January 3, 2000, a regular daily series of observations begins on July 9, 2001 at these maturities. 3-year OIS rates are available from March 3, 2004. 4 and 5-year OIS rates are available from July 19, 2005 and June 13, 2005, respectively.

Japan Ooka, Nagano, and Baba (2006) describe the growth in Japanese OIS markets during the years preceding 2006. Observations for 1, 2 and 3-month OIS rate data begin on March 15, 2002, but the daily series are sporadic. The first observations for the 6, 9 and 12-month OIS contracts is March 22, 2002, but the time series are also sporadic. Regular daily observations for the 1, 2, 3, 4, 5, 6, 9 and 12-month OIS rates are available from July 24, 2003. 7, 8, 10

and 11-month OIS rates are available from November 16, 2004. 15 and 21-month OIS rates are available from May 5, 2007. 18 and 24-month OIS rates are available from December 7, 2005. 3-year OIS rates are regularly available from November 19, 2007. 4-and 5-year OIS rates are available from August 6, 2009.

UK 1, 2, 3, 4, 5, 6, 7, 8, 9, 10, 11, 12, 18 and 24-month OIS rates are available from December 14, 2000. 15 and 21-month OIS rates are available from January 25, 2006. 3 and 4-year OIS rates are available from September 4, 2008. 5-year OIS rates are available from May 23, 2008.

US 1, 2, 3, 4, 5, 6, 9, 12 and 21-month OIS rates are available at a daily frequency from December 5, 2001. 15, 18 and 24-month OIS rates are available at from December 21, 2001. 7, 8, 10 and 11-month OIS rates are available from May 7, 2002. 3, 4 and 5-year OIS rates are available from February 14, 2002. 13, 14, 16, 17, 19, 20, 22 and 23-month OIS rates are only available from March 3, 2010 to June 14, 2011.

A.1.2 International Asset Portfolio Data

The below paragraphs describe the sources of international asset portfolio data from chapter 5.

Coordinated Portfolio Investor Survey (CPIS) This dataset is collated by the International Monetary Fund (IMF). The CPIS is a voluntary data collection exercise, in which participating economies provide data on its total holdings of portfolio investment securities. The survey quotes international asset holdings on December 31 for the 2001-2012 period. From 2013 onwards, the survey has been carried out twice a year: on June 30 and December 31. In chapter 5, I use the annual December surveys only from 2001 to 2014 — the last December survey available to date.

US Treasury International Capital (TIC) System The TIC reporting system the US government's source of data on capital flows into and out of the United States. The data is collected by the US Treasury and the Federal Reserve Bank of New York. In chapter 5, I use annual holdings data from the TIC system. This data is collated from surveys of issuers and holders of US and foreign securities held at the end of June each year. The first annual survey was conducted in 2002, and I use the survey data from 2002 to 2015. Prior to 2002, similar surveys were carried out, albeit less frequently. I use the data to study asset holdings by country of foreign holder, security type and maturity. However, more generally, the dataset permits analysis of currency of issuance, type of foreign holders and industry of asset issuer.

Two features of the TIC system data enhance its accuracy *vis-à-vis* the CPIS data. First, TIC surveys acquire information at the level of individual securities. Second, reporting for the TIC surveys is compulsory, and significant penalties can be imposed for failure to report.

A.2 Approximating Survey Forecasts

In chapters 3-5, I construct approximate survey forecasts of interest rates to compare estimates of interest rate expectations across financial market instruments, and term structure models. In this appendix, I describe how these survey forecast approximations are constructed.

I use data from the *Survey of Professional Forecasters* (SPF) at the Federal Reserve Bank of Philadelphia. The survey is published every quarter and reports forecasters' median expectations of the average 3-month T-Bill rate over a specified time period: the current quarter $\bar{i}_{t|t}^{3m,sur}$; and the first $\bar{i}_{t+1|t}^{3m,sur}$, second $\bar{i}_{t+2|t}^{3m,sur}$, third $\bar{i}_{t+3|t}^{3m,sur}$ and fourth $\bar{i}_{t+4|t}^{3m,sur}$ quarters subsequent to the current one, where t denotes the current quarter. All quantities are plotted on the survey submission deadline dates.

To construct a geometric approximation for the average expectation of the 3-month T-Bill rate over the 3-months following the deadline date, I construct an equally weighted geometric average of the median expectation of the 3-month rate for the current and the subsequent quarter. An equal weighting is made possible because the survey deadline date lies approximately halfway through the 'current' quarter. I use a geometric average to facilitate direct comparison with OIS contracts, which have a geometric structure.

To achieve this, I first use the survey expectation for the average 3-month T-Bill rate over the current quarter $\bar{i}_{t|t}^{3m,sur}$ and the realised average of the 3-month T-Bill rate in the current quarter up to the SPF deadline date $\bar{i}_t^{3m,real}$ to approximate the survey expectations for the average 3-month T-Bill rate over the remainder of the current quarter, denoted $\bar{i}_{t+|t}^{3m,sur}$. This is calculated from the following expression:

$$\bar{i}_{t|t}^{3m,sur} = \frac{1}{2}\bar{i}_t^{3m,real} + \frac{1}{2}\bar{i}_{t+|t}^{3m,sur}$$

For figure 3.1, I calculate the average survey expectation of the 3-month T-Bill rate over the 3, 6 and 12 months following the SPF deadline date. To calculate the average survey expectation of the 3-month T-Bill rate over the three months from the SPF deadline date t , $i_{t|t}^{3m,sur}$, I use the approximation:

$$i_{t|t}^{3m,sur} = \left[\left(1 + \frac{\bar{i}_{t+|t}^{3m,sur}}{100} \right)^{\frac{1}{2}} \times \left(1 + \frac{\bar{i}_{t+1|t}^{3m,sur}}{100} \right)^{\frac{1}{2}} - 1 \right] \times 100$$

where $i_{t|t}^{3m,sur}$, $\bar{i}_{t+|t}^{3m,sur}$ and $\bar{i}_{t+1|t}^{3m,sur}$ are all reported in percentage points.

The average expectation of the 3-month T-Bill rate over the six months following the deadline date t , $i_{t|t}^{6m,sur}$, is approximated using a similar geometric *weighted* average procedure: the expectation of the 3-month rate for the remainder of the current quarter and second quarter ahead are both given weights of 1/4; and the first-quarter-ahead expectation has weight 1/2. Mathematically, this is written as:

$$i_{t|t}^{6m,sur} = \left[\left(1 + \frac{\bar{i}_{t+|t}^{3m,sur}}{100} \right)^{\frac{1}{4}} \times \left(1 + \frac{\bar{i}_{t+1|t}^{3m,sur}}{100} \right)^{\frac{1}{2}} \times \left(1 + \frac{\bar{i}_{t+2|t}^{3m,sur}}{100} \right)^{\frac{1}{4}} - 1 \right] \times 100$$

The average expectation of the 3-month T-Bill rate over the year following the submission date t , $i_{t|t}^{1y,sur}$, is approximated by a geometric weighted average of the remainder of the current quarter

and first, second, third and fourth quarter ahead expectations, of the form:

$$i_{t|t}^{1y,sur} = \left[\left(1 + \frac{\bar{i}_{t+|t}^{3m,sur}}{100} \right)^{\frac{1}{8}} \times \left(1 + \frac{\bar{i}_{t+1|t}^{3m,sur}}{100} \right)^{\frac{1}{4}} \times \left(1 + \frac{\bar{i}_{t+2|t}^{3m,sur}}{100} \right)^{\frac{1}{4}} \right. \\ \left. \times \left(1 + \frac{\bar{i}_{t+3|t}^{3m,sur}}{100} \right)^{\frac{1}{4}} \times \left(1 + \frac{\bar{i}_{t+4|t}^{3m,sur}}{100} \right)^{\frac{1}{8}} - 1 \right] \times 100$$

In section 3.6.2, I calculate the average expectation of the 3-month T-Bill rate over the 1.5, 4.5, 7.5, 10.5 and 13.5 months following the deadline date. To do this, I construct weighted geometric averages of the median expectation of the 3-month rate for the current and subsequent quarters. The weighting, which facilitates direct comparison of the survey and risk-neutral yield-implied expectations, is made possible because the survey deadline date lies approximately halfway through the ‘current’ quarter. Mathematically, the average survey expectation of the 3-month T-Bill rate over the 1.5, 4.5, 7.5, 10.5 and 13.5 months from the deadline date t are given by:

$$i_{t|t}^{1.5m,sur} = \bar{i}_{t+|t}^{3m,sur}$$

$$i_{t|t}^{4.5m,sur} = \left[\left(1 + \frac{\bar{i}_{t+|t}^{3m,sur}}{100} \right)^{\frac{1}{3}} \times \left(1 + \frac{\bar{i}_{t+1|t}^{3m,sur}}{100} \right)^{\frac{2}{3}} - 1 \right] \times 100$$

$$i_{t|t}^{7.5m,sur} = \left[\left(1 + \frac{\bar{i}_{t+|t}^{3m,sur}}{100} \right)^{\frac{1}{5}} \times \left(1 + \frac{\bar{i}_{t+1|t}^{3m,sur}}{100} \right)^{\frac{2}{5}} \times \left(1 + \frac{\bar{i}_{t+2|t}^{3m,sur}}{100} \right)^{\frac{2}{5}} - 1 \right] \times 100$$

$$i_{t|t}^{10.5m,sur} = \left[\left(1 + \frac{\bar{i}_{t+|t}^{3m,sur}}{100} \right)^{\frac{1}{7}} \times \left(1 + \frac{\bar{i}_{t+1|t}^{3m,sur}}{100} \right)^{\frac{2}{7}} \times \left(1 + \frac{\bar{i}_{t+2|t}^{3m,sur}}{100} \right)^{\frac{2}{7}} \right. \\ \left. \times \left(1 + \frac{\bar{i}_{t+3|t}^{3m,sur}}{100} \right)^{\frac{2}{7}} - 1 \right] \times 100$$

$$i_{t|t}^{13.5m,sur} = \left[\left(1 + \frac{\bar{i}_{t+|t}^{3m,sur}}{100} \right)^{\frac{1}{9}} \times \left(1 + \frac{\bar{i}_{t+1|t}^{3m,sur}}{100} \right)^{\frac{2}{9}} \times \left(1 + \frac{\bar{i}_{t+2|t}^{3m,sur}}{100} \right)^{\frac{2}{9}} \right. \\ \left. \times \left(1 + \frac{\bar{i}_{t+3|t}^{3m,sur}}{100} \right)^{\frac{2}{9}} \times \left(1 + \frac{\bar{i}_{t+4|t}^{3m,sur}}{100} \right)^{\frac{2}{9}} - 1 \right] \times 100$$

Appendix B

Appendix to Chapter 3

B.1 Baseline Gaussian Affine Dynamic Term Structure Model

B.1.1 Bond Pricing Using the Risk-Adjusted Probability Measure \mathbb{Q}

To guarantee the existence of a risk-adjusted probability measure \mathbb{Q} , under which the bonds are priced, no-arbitrage is imposed (Harrison and Kreps, 1979). The risk-adjusted probability measure \mathbb{Q} is defined such that the price V_t of any asset that does not pay any dividends at time $t+1$ satisfies $V_t = \mathbb{E}_t^{\mathbb{Q}} [\exp(-i_t)V_{t+1}]$, where the expectation $\mathbb{E}_t^{\mathbb{Q}}$ is taken under the \mathbb{Q} probability measure. Thus, with no-arbitrage, the price of an n -day zero-coupon bond must satisfy the following relation:

$$P_{t,n} = \mathbb{E}_t^{\mathbb{Q}} [\exp(-i_t)P_{t+1,n-1}] \quad (12)$$

Using this, it is possible to show that the nominal bond price is an exponentially affine function of the pricing factors:

$$P_{t,n} = \exp(\mathcal{A}_n + \mathcal{B}_n \mathbf{x}_t) \quad (13)$$

such that the corresponding continuously compounded yield $y_{t,n}$ is affine in the pricing factors:

$$y_{t,n} = -\frac{1}{n} \ln(P_{t,n}) = A_n + B_n \mathbf{x}_t \quad (\text{B.1})$$

where $A_n \equiv -\frac{1}{n} \mathcal{A}_n(\delta_0, \delta_1, \boldsymbol{\mu}^{\mathbb{Q}}, \boldsymbol{\Phi}^{\mathbb{Q}}, \boldsymbol{\Sigma}; \mathcal{A}_{n-1}, \mathcal{B}_{n-1})$ and $B_n \equiv -\frac{1}{n} \mathcal{B}_n(\delta_1, \boldsymbol{\Phi}^{\mathbb{Q}}; \mathcal{B}_{n-1})$.

To attain recursive expressions for \mathcal{A}_n and \mathcal{B}_n :

$$\begin{aligned} \mathcal{A}_n + \mathcal{B}_n \mathbf{x}_t &= \ln P_{t,n} \\ &= \ln \mathbb{E}_t^{\mathbb{Q}} [\exp(-i_t)P_{t+1,n-1}] \\ &= \ln \mathbb{E}_t^{\mathbb{Q}} [\exp(-i_t + \mathcal{A}_{n-1} + \mathcal{B}_{n-1} \mathbf{x}_{t+1})] \\ &= \ln \mathbb{E}_t^{\mathbb{Q}} \left[\exp \left(-\delta_0 - \delta_1' \mathbf{x}_t + \mathcal{A}_{n-1} + \mathcal{B}_{n-1} \left[\boldsymbol{\mu}^{\mathbb{Q}} + \boldsymbol{\Phi}^{\mathbb{Q}} \mathbf{x}_t + \boldsymbol{\Sigma} \boldsymbol{\varepsilon}_{t+1}^{\mathbb{Q}} \right] \right) \right] \\ &= -(\delta_0 + \delta_1' \mathbf{x}_t) + \mathcal{A}_{n-1} + \mathcal{B}_{n-1} \left[\boldsymbol{\mu}^{\mathbb{Q}} + \boldsymbol{\Phi}^{\mathbb{Q}} \mathbf{x}_t \right] + \ln \mathbb{E}_t^{\mathbb{Q}} \left[\exp \left(\mathcal{B}_{n-1} \boldsymbol{\Sigma} \boldsymbol{\varepsilon}_{t+1}^{\mathbb{Q}} \right) \right] \\ &= -(\delta_0 + \delta_1' \mathbf{x}_t) + \mathcal{A}_{n-1} + \mathcal{B}_{n-1} \left[\boldsymbol{\mu}^{\mathbb{Q}} + \boldsymbol{\Phi}^{\mathbb{Q}} \mathbf{x}_t \right] + \frac{1}{2} \mathcal{B}_{n-1} \boldsymbol{\Sigma} \boldsymbol{\Sigma}' \mathcal{B}_{n-1}' \\ &= \left\{ -\delta_0 + \mathcal{A}_{n-1} + \frac{1}{2} \mathcal{B}_{n-1} \boldsymbol{\Sigma} \boldsymbol{\Sigma}' \mathcal{B}_{n-1}' + \mathcal{B}_{n-1} \boldsymbol{\mu}^{\mathbb{Q}} \right\} + \left\{ -\delta_1' + \mathcal{B}_{n-1} \boldsymbol{\Phi}^{\mathbb{Q}} \right\} \mathbf{x}_t \end{aligned}$$

using (13) in the third line, (3.2) and (3.5) in the fourth line, and using the property of the

log-normal distribution in conjunction with the fact that $\varepsilon_{t+1}^{\mathbb{Q}} | \mathbf{x}_t \sim \mathcal{N}(\mathbf{0}_K, \mathbf{I}_K)$ to write the expression $\ln \mathbb{E}_t^{\mathbb{Q}} \left[\exp \left(\mathcal{B}_{n-1} \Sigma \varepsilon_{t+1}^{\mathbb{Q}} \right) \right]$ as $\frac{1}{2} \mathcal{B}_{n-1} \Sigma \Sigma' \mathcal{B}_{n-1}'$ in the sixth line.

By the method of undetermined coefficients, the recursive definitions for the scalar $\mathcal{A}_n \equiv \mathcal{A}_n(\delta_0, \delta_1, \boldsymbol{\mu}^{\mathbb{Q}}, \boldsymbol{\Phi}^{\mathbb{Q}}, \Sigma; \mathcal{A}_{n-1}, \mathcal{B}_{n-1})$ and the $1 \times K$ vector $\mathcal{B}_n \equiv \mathcal{B}_n(\delta_1, \boldsymbol{\Phi}^{\mathbb{Q}}; \mathcal{B}_{n-1})$ follow from the final line:

$$\mathcal{A}_n = -\delta_0 + \mathcal{A}_{n-1} + \frac{1}{2} \mathcal{B}_{n-1} \Sigma \Sigma' \mathcal{B}_{n-1}' + \mathcal{B}_{n-1} \boldsymbol{\mu}^{\mathbb{Q}} \quad (\text{B.2})$$

$$\mathcal{B}_n = -\delta_1' + \mathcal{B}_{n-1} \boldsymbol{\Phi}^{\mathbb{Q}} \quad (\text{B.3})$$

with initial values $\mathcal{A}_0 = 0$ and $\mathcal{B}_0 = \mathbf{0}_K'$, where $\mathbf{0}_K$ is a $K \times 1$ vector of zeros.

B.1.2 Bond Pricing Using the Pricing Kernel and the Actual Probability Measure \mathbb{P}

Under the actual \mathbb{P} probability measure, the bond price is given by equation (3.7):

$$P_{t,n} = \mathbb{E}_t [M_{t+1} P_{t+1,n-1}]$$

where this expectation is taken under the \mathbb{P} measure.

Using this, it is also possible to show that the nominal bond price is an exponentially affine function of the pricing factors, as in (13). To attain recursive expressions for \mathcal{A}_n and \mathcal{B}_n :

$$\begin{aligned} \mathcal{A}_n + \mathcal{B}_n \mathbf{x}_t &= \ln P_{t,n} \\ &= \ln \mathbb{E}_t [M_{t+1} P_{t+1,n-1}] \\ &= \ln \mathbb{E}_t \left[\exp \left(-i_t - \frac{1}{2} \boldsymbol{\lambda}_t' \boldsymbol{\lambda}_t - \boldsymbol{\lambda}_t' \varepsilon_{t+1} + \mathcal{A}_{n-1} + \mathcal{B}_{n-1} \mathbf{x}_{t+1} \right) \right] \\ &= \ln \mathbb{E}_t \left[\exp \left(-\delta_0 - \delta_1' \mathbf{x}_t - \frac{1}{2} (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t)' (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t) \right. \right. \\ &\quad \left. \left. - (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t)' \varepsilon_{t+1} + \mathcal{A}_{n-1} + \mathcal{B}_{n-1} (\boldsymbol{\mu} + \boldsymbol{\Phi} \mathbf{x}_t + \Sigma \varepsilon_{t+1}) \right) \right] \\ &= -\delta_0 - \delta_1' \mathbf{x}_t - \frac{1}{2} (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t)' (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t) + \mathcal{A}_{n-1} + \mathcal{B}_{n-1} (\boldsymbol{\mu} + \boldsymbol{\Phi} \mathbf{x}_t) \\ &\quad + \ln \mathbb{E}_t \left[\exp \left((- (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t)' + \mathcal{B}_{n-1} \Sigma) \varepsilon_{t+1} \right) \right] \\ &= -\delta_0 - \delta_1' \mathbf{x}_t - \frac{1}{2} (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t)' (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t) + \mathcal{A}_{n-1} + \mathcal{B}_{n-1} (\boldsymbol{\mu} + \boldsymbol{\Phi} \mathbf{x}_t) \\ &\quad + \frac{1}{2} (- (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t)' + \mathcal{B}_{n-1} \Sigma) (- (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t)' + \mathcal{B}_{n-1} \Sigma)' \\ &= -\delta_0 - \delta_1' \mathbf{x}_t + \mathcal{A}_{n-1} - \mathcal{B}_{n-1} \Sigma (\boldsymbol{\lambda}_0 + \boldsymbol{\Lambda}_1 \mathbf{x}_t) \\ &\quad + \frac{1}{2} \mathcal{B}_{n-1} \Sigma \Sigma' \mathcal{B}_{n-1}' + \mathcal{B}_{n-1} (\boldsymbol{\mu} + \boldsymbol{\Phi} \mathbf{x}_t) \end{aligned}$$

using (3.3) and (13) in the third line, and (3.1), (3.2) and (3.4) in the fourth line.

By the method of undetermined coefficients, the recursive definitions for the scalar \mathcal{A}_n and the $1 \times K$ vector \mathcal{B}_n follow from the final line:

$$\mathcal{A}_n = -\delta_0 + \mathcal{A}_{n-1} + \frac{1}{2} \mathcal{B}_{n-1} \Sigma \Sigma' \mathcal{B}_{n-1}' + \mathcal{B}_{n-1} (\boldsymbol{\mu} - \Sigma \boldsymbol{\lambda}_0) \quad (\text{B.4})$$

$$\mathcal{B}_n = -\delta_1' + \mathcal{B}_{n-1} (\boldsymbol{\Phi} - \Sigma \boldsymbol{\Lambda}_1) \quad (\text{B.5})$$

with initial values $\mathcal{A}_0 = 0$ and $\mathcal{B}_0 = \mathbf{0}_K'$, where $\mathbf{0}_K'$ is a $K \times 1$ vector of zeros.

Comparing (B.2) and (B.3) with (B.4) and (B.5) yields the relationship between \mathbb{P} and \mathbb{Q} parameters:

$$\boldsymbol{\mu}^{\mathbb{Q}} = \boldsymbol{\mu} - \boldsymbol{\Sigma}\boldsymbol{\lambda}_0, \quad \boldsymbol{\Phi}^{\mathbb{Q}} = \boldsymbol{\Phi} - \boldsymbol{\Sigma}\boldsymbol{\Lambda}_1.$$

B.1.3 Risk-Neutral Yields

The risk-neutral yield on an n -day bond reflects the yield that would prevail if investors were risk-neutral. That is, the risk-neutral yield corresponds to that which would arise under the actual probability measure \mathbb{P} .

The risk-neutral bond price $\tilde{P}_{t,n}$ is of the form:

$$\tilde{P}_{t,n} = \mathbb{E}_t \left[\exp(-i_t) \tilde{P}_{t+1,n-1} \right] \quad (\text{B.6})$$

and can be shown to be an exponentially affine function of the pricing factors:

$$\tilde{P}_{t,n} = \exp(\mathcal{A}_n + \mathcal{B}_n \mathbf{x}_t) \quad (\text{B.7})$$

where $\mathcal{A}_n \equiv \mathcal{A}_n(\delta_0, \boldsymbol{\delta}_1, \boldsymbol{\mu}, \boldsymbol{\Phi}, \boldsymbol{\Sigma}; \mathcal{A}_{n-1}, \mathcal{B}_{n-1})$ and $\mathcal{B}_n \equiv \mathcal{B}_n(\boldsymbol{\delta}_1, \boldsymbol{\Phi}; \mathcal{B}_{n-1})$. Thus, the risk-neutral yield, $\tilde{y}_{t,n} = -\frac{1}{n} \ln \tilde{P}_{t,n}$, is affine in the pricing factors:

$$\tilde{y}_{t,n} = \tilde{A}_n + \tilde{B}_n \mathbf{x}_t \quad (\text{B.8})$$

where $\tilde{A}_n = -\frac{1}{n} \mathcal{A}_n(\delta_0, \boldsymbol{\delta}_1, \boldsymbol{\mu}, \boldsymbol{\Phi}, \boldsymbol{\Sigma}; \mathcal{A}_{n-1}, \mathcal{B}_{n-1})$ and $\tilde{B}_n = -\frac{1}{n} \mathcal{B}_n(\boldsymbol{\delta}_1, \boldsymbol{\Phi}; \mathcal{B}_{n-1})$.

To attain the recursive expressions for \tilde{A}_n and \tilde{B}_n , note that from equation (B.7):

$$\begin{aligned} \tilde{y}_{t,n} &= -\frac{1}{n} \ln \mathbb{E}_t [\exp \{-i_t + \mathcal{A}_{n-1} + \mathcal{B}_{n-1} \mathbf{x}_{t+1}\}] \\ \tilde{A}_n + \tilde{B}_n \mathbf{x}_t &= -\frac{1}{n} \ln \mathbb{E}_t [\exp \{-(\delta_0 + \boldsymbol{\delta}'_1 \mathbf{x}_t) + \mathcal{A}_{n-1} + \mathcal{B}_{n-1} [\boldsymbol{\mu} + \boldsymbol{\Phi} \mathbf{x}_t + \boldsymbol{\Sigma} \boldsymbol{\varepsilon}_{t+1}]\}] \\ &= -\frac{1}{n} \left[\left\{ -\delta_0 + \mathcal{A}_{n-1} + \frac{1}{2} \mathcal{B}_{n-1} \boldsymbol{\Sigma} \boldsymbol{\Sigma}' \mathcal{B}_{n-1}' + \mathcal{B}_{n-1} \boldsymbol{\mu} \right\} + \{ -\boldsymbol{\delta}'_1 + \mathcal{B}_{n-1} \boldsymbol{\Phi} \} \mathbf{x}_t \right] \end{aligned}$$

using (B.6) and (B.7) in the first line, and (3.2) and (3.1) in the second line. The expectation is taken under the actual probability measure \mathbb{P} . By the method of undetermined coefficients, it follows that:

$$\tilde{y}_{t,n} = \tilde{A}_n + \tilde{B}_n \mathbf{x}_t$$

where $\tilde{A}_n = -\frac{1}{n} \mathcal{A}_n(\delta_0, \boldsymbol{\delta}_1, \boldsymbol{\mu}, \boldsymbol{\Phi}, \boldsymbol{\Sigma}; \mathcal{A}_{n-1}, \mathcal{B}_{n-1})$ and $\tilde{B}_n = -\frac{1}{n} \mathcal{B}_n(\boldsymbol{\delta}_1, \boldsymbol{\Phi}; \mathcal{B}_{n-1})$.

B.2 Overnight Indexed Swap Augmentation

To calculate the loadings for the OIS observation equation, first note that the *annualised* floating leg of a j -day OIS contract, with trade day t and starting day $t+1$, is given by:

$$i_{t,t+j}^{flt} = \left(\left[\prod_{k=1}^j (1 + \gamma_{t+k} i_{t+k}) \right] - 1 \right) \times \frac{T_{yr}}{j}$$

where γ_{t+k} is an accrual factor, which is set to $1/T_{yr}$ for all time periods, and $T_{yr} = 252$ is the number of trading days in a year.¹ i_t is the one-period short-term floating interest rate (3.2) used as the reference rate for the swap — the effective federal funds rate. Rearranging this, taking logs, and using a first-order Taylor approximation around $x = 0$ such that $\ln(1+x) \approx x$, yields:

$$i_{t,t+k}^{flt} \approx \left[\sum_{k=1}^j \gamma_{t+k} i_{t+k} \right] \times \frac{T_{yr}}{j} \quad (\text{B.9})$$

Therefore, under the expectations hypothesis, the OIS rate $i_{t,t+j}^{ois}$ will be:

$$i_{t,t+j}^{ois} = \mathbb{E}_t \left[\sum_{k=1}^j \gamma_{t+k} i_{t+k} \right] \times \frac{T_{yr}}{j} \quad (\text{B.10})$$

For a one-period OIS contract, $j = 1$:

$$\begin{aligned} i_{t,t+1}^{ois} &= \mathbb{E}_t [\gamma (\delta_0 + \delta'_1 \mathbf{x}_{t+1})] \times T_{yr} \\ &= \mathbb{E}_t [\gamma (\delta_0 + \delta'_1 [\boldsymbol{\mu} + \boldsymbol{\Phi} \mathbf{x}_t + \boldsymbol{\Sigma} \boldsymbol{\varepsilon}_{t+1}])] \times T_{yr} \\ &= (\delta_0 + \delta'_1 \boldsymbol{\mu} + \delta'_1 \boldsymbol{\Phi} \mathbf{x}_t) \end{aligned}$$

where $\gamma \equiv 1/T_{yr}$, and noting that $\mathbb{E}_t [\boldsymbol{\varepsilon}_{t+1}] = \mathbf{0}_K$ in the second line.

For a two period OIS contract, $j = 2$, the expectations hypothesis requires that:

$$\begin{aligned} i_{t,t+2}^{ois} &= \mathbb{E}_t [\gamma (\delta_0 + \delta'_1 \boldsymbol{\mu} + \delta'_1 \boldsymbol{\Phi} \mathbf{x}_t) + \gamma (\delta_0 + \delta'_1 \boldsymbol{\mu} + \delta'_1 \boldsymbol{\Phi} \mathbf{x}_{t+1})] \times (T_{yr}/2) \\ &= \mathbb{E}_t [\gamma (\delta_0 + \delta'_1 \boldsymbol{\mu} + \delta'_1 \boldsymbol{\Phi} \mathbf{x}_t) + \gamma (\delta_0 + \delta'_1 \boldsymbol{\mu} + \delta'_1 \boldsymbol{\Phi} [\boldsymbol{\mu} + \boldsymbol{\Phi} \mathbf{x}_t + \boldsymbol{\Sigma} \boldsymbol{\varepsilon}_{t+1}])] \times (T_{yr}/2) \\ &= \frac{1}{2} (2\delta_0 + 2\delta'_1 \boldsymbol{\mu} + \delta'_1 \boldsymbol{\Phi} \boldsymbol{\mu} + \delta'_1 \boldsymbol{\Phi} \mathbf{x}_t + \delta'_1 \boldsymbol{\Phi}^2 \mathbf{x}_t) \end{aligned}$$

For a three period OIS contract, $j = 3$, the same steps as above yield the following expression:

$$i_{t,t+3}^{ois} = \frac{1}{3} (3\delta_0 + 3\delta'_1 \boldsymbol{\mu} + 2\delta'_1 \boldsymbol{\Phi} \boldsymbol{\mu} + \delta'_1 \boldsymbol{\Phi}^2 \boldsymbol{\mu} + \delta'_1 \boldsymbol{\Phi} \mathbf{x}_t + \delta'_1 \boldsymbol{\Phi}^2 \mathbf{x}_t + \delta'_1 \boldsymbol{\Phi}^3 \mathbf{x}_t)$$

The continued iteration can be summarised by the following expressions:

$$i_{t,t+j}^{ois} = A_j^{ois} + B_j^{ois} \mathbf{x}_t \quad (18)$$

where $A_j^{ois} \equiv \frac{1}{j} \mathcal{A}_j^{ois} (\delta_0, \delta_1, \boldsymbol{\mu}, \boldsymbol{\Phi}, \boldsymbol{\Sigma}; \mathcal{A}_{j-1}^{ois}, \mathcal{B}_{j-1}^{ois})$ and $B_j^{ois} = \frac{1}{j} \mathcal{B}_j^{ois} (\delta_1, \boldsymbol{\Phi}; \mathcal{B}_{j-1}^{ois})$ are recursively defined as:

$$\begin{aligned} \mathcal{A}_j^{ois} &= \delta_0 + \delta'_1 \boldsymbol{\mu} + \mathcal{A}_{j-1}^{ois} + \mathcal{B}_{j-1}^{ois} \boldsymbol{\mu} \\ \mathcal{B}_j^{ois} &= \delta'_1 \boldsymbol{\Phi} + \mathcal{B}_{j-1}^{ois} \boldsymbol{\Phi} \end{aligned}$$

where $\mathcal{A}_0^{ois} = 0$ and $\mathcal{B}_0^{ois} = \mathbf{0}'_K$, where $\mathbf{0}_K$ is a $K \times 1$ vector of zeros.

¹For the term structure model, the accrual and annualisation factors use the convention that there are 252 business trading days in a year, as opposed to the market quoting convention of 360 days used in section 3.2. Given that daily yield data is only available on 252 days per year, I adopt this convention to ensure that the *horizon* for each OIS rate corresponds to their actual maturity date and that of a corresponding maturity zero-coupon bond. This convention is also adopted for daily frequency term structure estimation by, amongst others, Bauer and Rudebusch (2014).

I attain an affine expression for OIS rates (3.14), with loadings that are additive and recursive, because I use a first-order Taylor approximation of $i_{t,t+j}^{flt}$ in (B.9). This ensures that OIS rates can be included in the GADTSM in a similar manner to bond yields, which are also affine in the pricing factors \mathbf{x}_t . In reality, because the floating leg of an OIS contract is compounded, a Jensen's inequality term would be expected in the OIS pricing expression, representing a term premium in OIS rates. The first-order Taylor approximation of $i_{t,t+j}^{ois}$ (B.9) prevents a Jensen's inequality term from arising, and considerably simplifies the expressions for loadings, \mathcal{A}_j^{ois} and \mathcal{B}_j^{ois} , ensuring that they are additive and recursive. I circumvent this potential problem by only using OIS rates which I demonstrate have statistically insignificant *ex post* excess returns, such that any Jensen's inequality term should be negligible. Furthermore, I continue to admit measurement error in OIS rates in the Kalman filter setup (3.15) through \mathbf{u}_t^{ois} .

B.3 Estimation Procedure

To identify the unaugmented model described in section 3.3, I use the normalisation scheme proposed by Joslin et al. (2011). The Joslin et al. (2011) normalisation fosters faster convergence to the global optimum of the model's likelihood function than other identification schemes for two reasons. First, this normalisation allows for the (near) separation of the \mathbb{P} and \mathbb{Q} probability measure likelihood functions, the product of which comprises the overall model likelihood function. Moreover, the Joslin et al. (2011) normalisation reduces the dimensionality of the parameter space. In the baseline, unaugmented model, the parameters governing bond pricing are:

$$\Theta = \left\{ \delta_0, \delta_1, \boldsymbol{\mu}^{\mathbb{Q}}, \boldsymbol{\Phi}^{\mathbb{Q}}, \boldsymbol{\Sigma} \right\}$$

The Joslin et al. (2011) normalisation scheme uniquely maps these parameters to a smaller set:

$$\left\{ i_{\infty}^{\mathbb{Q}}, \boldsymbol{\lambda}^{\mathbb{Q}}, \boldsymbol{\Sigma} \right\}$$

where: (i) $i_{\infty}^{\mathbb{Q}}$ is the risk-neutral expectation of the long-run short-term nominal interest rate; (ii) $\boldsymbol{\lambda}^{\mathbb{Q}}$ is a $K \times 1$ vector of the eigenvalues of $\boldsymbol{\Phi}^{\mathbb{Q}}$; and (iii) $\boldsymbol{\Sigma}$ is a lower triangular matrix with positive diagonal entries.

B.3.1 OLS/ML Estimation of the Baseline, Unaugmented GADTSM

Assuming that K portfolios of bonds are priced without error, then the Joslin et al. (2011) normalisation permits the complete separation of the \mathbb{P} and \mathbb{Q} likelihood functions. In this study, as in many others, I use the first K principal components of the observed bond yields as the set of K portfolios that are priced perfectly (e.g. Joslin et al., 2011). Defining these portfolios $\mathcal{P}_t \equiv W y_t = W y_t^{obs} \equiv \mathcal{P}_t^{obs}$, where W is the principal component weighting matrix and y_t^{obs} is the vector of observed yields, then Joslin et al. (2011) show that the likelihood function for the unaugmented model laid out in section 3.3.1 is:

$$\mathcal{L} \left(y_t^{obs} | y_{t-1}^{obs}; \Theta \right) = \mathcal{L} \left(y_t^{obs} | \mathcal{P}_t; \boldsymbol{\lambda}^{\mathbb{Q}}, i_{\infty}^{\mathbb{Q}}, \boldsymbol{\Sigma}, \sigma_u \right) \times \mathcal{L} \left(\mathcal{P}_t | \mathcal{P}_{t-1}; \boldsymbol{\mu}, \boldsymbol{\Phi}, \boldsymbol{\Sigma} \right)$$

where σ_u is the standard deviation of the measurement error of the N observed yields.

This normalisation admits a two-stage estimation process. First, the parameters $\{\boldsymbol{\mu}, \boldsymbol{\Phi}\}$ are

directly estimable by running OLS on the VAR in equation (3.1), where $\mathbf{x}_t \equiv \mathcal{P}_t$. Moreover, this provides initial values for the maximum likelihood estimation of the lower triangular elements of the matrix $\mathbf{\Sigma}$. Second, taking $\{\hat{\boldsymbol{\mu}}, \hat{\mathbf{\Phi}}\}$ as given, the parameters $\{i_{\infty}^{\mathbb{Q}}, \boldsymbol{\lambda}^{\mathbb{Q}}, \mathbf{\Sigma}, \sigma_u\}$ can be estimated by maximum likelihood.

B.3.2 Bias-Corrected Estimation

To estimate the bias-corrected decomposition, I rely entirely on the methodology of Bauer et al. (2012, Section 4). The MATLAB code for this is available here: faculty.chicagobooth.edu/jing.wu/research/zip/brw_table1.zip.

B.3.3 Survey-Augmentation

To augment the model with survey expectations of future interest rates, I employ Kalman filter-based maximum likelihood estimation. This estimation methodology, using survey expectations, draws most directly on Guimarães (2014).

Like Guimarães (2014), I use survey expectations from the *Survey of Professional Forecasters* at the Federal Reserve Bank of Philadelphia. I use forecasts for the 3 month T-Bill 1, 2, 3 and 4 quarters ahead, available at a quarterly frequency. I augment the model with the survey expectations on the survey submission deadline day.²

The survey-augmented Kalman filter has a similar form to the OIS-augmented setup presented in section 3.3. The transition equation of the Kalman filter is (3.1), the vector autoregression for the latent pricing factors under the actual \mathbb{P} probability measure.

On days when the survey forecasts are *not* observed, the observation equation is given by (3.13). As with the OIS-augmented model, I maintain a homoskedastic form for the yield measurement error.

On days when the S survey forecasts, $s = s_1, s_2, \dots, s_S$, are observed, the observation equation is:

$$\begin{bmatrix} \mathbf{y}_t \\ \mathbf{i}_t^{sur} \end{bmatrix} = \begin{bmatrix} \mathbf{A} \\ \mathbf{A}^{sur} \end{bmatrix} + \begin{bmatrix} \mathbf{B} \\ \mathbf{B}^{sur} \end{bmatrix} \mathbf{x}_t + \begin{bmatrix} \mathbf{\Sigma}_Y & \mathbf{0}_{N \times S} \\ \mathbf{0}_{S \times N} & \mathbf{\Sigma}_S \end{bmatrix} \begin{bmatrix} \mathbf{u}_t \\ \mathbf{u}_t^{sur} \end{bmatrix} \quad (\text{B.11})$$

where, in addition to the definitions of \mathbf{y}_t , \mathbf{A} , \mathbf{B} , $\mathbf{\Sigma}_Y$ and \mathbf{u}_t above, $\mathbf{i}_t^{sur} = [i_{t,s_1}^{sur}, \dots, i_{t,s_S}^{sur}]'$; $\mathbf{A}^{sur} = [A_{s_1}^{sur}, \dots, A_{s_S}^{sur}]'$; $\mathbf{B}^{sur} = [B_{s_1}^{sur}, \dots, B_{s_S}^{sur}]'$; $\mathbf{0}_{S \times N}$ and $\mathbf{0}_{N \times S}$ denote $S \times N$ and $N \times S$ matrices of zeros respectively; and $\mathbf{u}_t^{sur} \sim \mathcal{N}(\mathbf{0}_S, \mathbf{I}_S)$ denotes the survey measurement error, where $\mathbf{0}_S$ is an S -vector of zeros and \mathbf{I}_S is an $S \times S$ identity matrix. As with the yield measurement error, I impose a homoskedastic form for the survey measurement error, such that $\mathbf{\Sigma}_S$ is a $S \times S$ diagonal matrix with common diagonal element σ_s , the standard deviation of the survey measurement error. Appendix C of Guimarães (2014) presents the functional forms for A_s^{sur} and B_s^{sur} , which account for the arithmetic nature of survey expectations.

As with the OIS-augmented model, I estimate the survey-augmented model by using the OLS/ML parameter estimates as initial values for the Kalman filter.

²For survey submission dates that are not business days, I augment the model with survey data on the preceding business day.

Table B.1: GADTSM Fit: Root Mean Square Error (RMSE) of the Fitted Yields *vis-à-vis* the Actual Yields

Sample: January 2002 to December 2016						
Maturity	OLS/ML	BC	Survey	2-OIS	3-OIS	4-OIS
3-Months	0.0979	0.0983	0.1028	0.1000	0.1091	0.1076
6-Months	0.0516	0.0513	0.0530	0.0524	0.0489	0.0534
1-Year	0.0714	0.0717	0.0775	0.0774	0.0777	0.0739
18-Months	0.0567	0.0564	0.0591	0.0598	0.0598	0.0605
2-Years	0.0403	0.0396	0.0395	0.0404	0.0396	0.0435
30-Months	0.0240	0.0234	0.0228	0.0237	0.0222	0.0265
3-Years	0.0161	0.0159	0.0181	0.0182	0.0177	0.0179
42-Months	0.0223	0.0223	0.0256	0.0249	0.0262	0.0237
4-Years	0.0313	0.0311	0.0339	0.0328	0.0349	0.0330
54-Months	0.0378	0.0374	0.0393	0.0380	0.0405	0.0401
5-Years	0.0410	0.0403	0.0414	0.0400	0.0425	0.0440
7-Years	0.0273	0.0263	0.0267	0.0265	0.0249	0.0333
10-Years	0.0638	0.0629	0.0609	0.0559	0.0608	0.0558
Average	0.0499	0.0497	0.0517	0.0507	0.0526	0.0525

Note: RMSE of the fitted yields from each of the six three-factor GADTSMs, computed by comparing the model-implied fitted yield to the actual yield on each day. All figures are expressed in annualised percentage points. The six GADTSMs are: (i) the unaugmented model estimated by OLS and maximum likelihood (OLS/ML); (ii) the bias-corrected model (BC); (iii) the survey-augmented model (Survey); (iv) the 2-OIS-augmented model (2-OIS); (v) the 3-OIS-augmented model (3-OIS); and (vi) the 4-OIS-augmented model (4-OIS).

B.3.4 OIS-Augmentation of the GADTSM and Kalman Filtering

When the Kalman filter is used, the assumption that K portfolios of yields are observed without error is no longer made. Instead, all yields (and portfolios thereof) can be observed with error. Consequently, the exact separation of the likelihood function described in section B.3.1 is no longer applicable. However, the parameter estimates attained from OLS/ML estimation of the unaugmented model do provide initial values for the Kalman filter-based optimisation routine.³ Doing so, ensures that computational time is reasonably fast.

B.4 Term Structure Results

In this section, I present additional results from the GADTSM estimation.

B.4.1 Additional Results for the Three-Factor Specification

Model-Implied Fitted Yields

Table B.1 presents the root mean square error (RMSE) for the fitted yields from each of the term structure models. The RMSE is presented for each maturity, and the average over all maturities, for the sample period January 2002 to December 2016. The fit, compared to the actual yield, is broadly similar across all six models. Specifically, the average RMSE for each of the models at all thirteen maturities is around 5 basis points, and differs by no more than 0.29 basis points across models. Thus, all models fit actual bond yields similarly well.

³Guimarães (2014) follows similar steps to estimate a survey-augmented GADTSM using the Joslin et al. (2011) normalisation scheme.

Model-Implied Fitted OIS Rates

Table B.2 presents the RMSE for the fitted OIS rates from each of the OIS-augmented term structure models. The RMSE is presented for each maturity for the sample period January 2002 to December 2016. The results demonstrate that the 4-OIS-augmented model provides superior estimates of the 6, 12 and 24-month OIS rates, while the 2-OIS-augmented model provides marginally superior estimates of the 3-month OIS rate. At the 6 and 12-month horizons, the 3-OIS-augmented model is only marginally inferior to the 4-OIS-augmented model. However, at the 2-year horizon, the 4-OIS-augmented model provides a substantial improvement in fit *vis-à-vis* the 3 and 2-OIS-augmented models. The 2-OIS-augmented provides the highest RMSE estimates of 1 and 2-year OIS rates.

Table B.2: GADTSM Fit: Root Mean Square Error (RMSE) of Fitted OIS Rates *vis-à-vis* the Actual OIS Rates

Sample: January 2002 to December 2016			
Maturity	2-OIS	3-OIS	4-OIS
3-Months	0.1183	0.1206	0.1305
6-Months	0.0924	0.1088	0.0861
1-Year	0.1585	0.0926	0.0831
2-Year	0.5306	0.2634	0.0985

Note: RMSE of the fitted OIS rates from each of the three OIS-augmented GADTSMs, computed by comparing the model-implied fitted OIS rate to the actual OIS rate on each day. All figures are expressed in annualised percentage points. The three GADTSMs are: (i) the 2-OIS-augmented model (2-OIS); (ii) the 3-OIS-augmented model (3-OIS); and (iii) the 4-OIS-augmented model (4-OIS).

Estimated Pricing Factors and Principal Components

Table B.3 presents summary statistics for the estimated principal components of the actual bond yields and the estimated pricing factors from the 4, 3 and 2-OIS-augmented models for the sample period January 2002 to December 2016. The results demonstrate that the principal components and estimated pricing factors evolve similarly, implying that OIS rates do not include any additional information, over and above that in bond yields, of value to the fitting of actual yields. In particular, the summary statistics of the estimated principal components and the estimated pricing factors from the 4-OIS-augmented models are similar.

Moreover, table B.3 further demonstrates that the inclusion of different maturities of OIS rate in the term structure model does not appreciably alter estimates of actual bond yields. The summary statistics of the estimated pricing factors from the 4, 3 and 2-OIS-augmented models are all similar. Augmentation of GADTSMs with OIS rates only influences estimated parameters under the actual probability measure \mathbb{P} and thus risk-neutral yields.

Stability of Estimates

In order to apply the term structure model to policy analysis, it is desirable for estimates of fitted yields, risk-neutral yields and term premia on any given date to be stable across different sample

Table B.3: Estimated Principal Components and Estimated Pricing Factors: Summary Statistics

Summary Statistics	1st Factor	2nd Factor	3rd Factor
Estimated Principal Components			
Mean	0.0725	0.0326	0.0075
Variance	0.0026	0.0001	0.0000
Skewness	0.7771	0.3839	−0.2603
Kurtosis	2.2918	2.2017	2.2441
4-OIS: Estimated Pricing Factors			
Mean	0.0725	0.0326	0.0076
Variance	0.0025	0.0001	0.0000
Skewness	0.7712	0.4043	−0.2515
Kurtosis	2.2792	2.2136	2.2732
3-OIS: Estimated Pricing Factors			
Mean	0.0725	0.0326	0.0075
Variance	0.0025	0.0001	0.0000
Skewness	0.7756	0.3909	−0.2542
Kurtosis	2.2877	2.1886	2.2772
2-OIS: Estimated Pricing Factors			
Mean	0.0725	0.0326	0.0075
Variance	0.0025	0.0001	0.0000
Skewness	0.7756	0.3797	−0.2812
Kurtosis	2.2893	2.1930	2.2152

Note: Summary statistics for the first three estimated principal components from actual yield data and the estimated pricing factors from the 4, 3 and 2-OIS-augmented models. All statistics are reported to four decimal places.

periods. That is, one would like a model that provides similar estimates of interest rate expectations on a given date regardless of the sample period used. In principle, OIS-augmentation can be helpful in this regard. By augmenting the model with additional information about interest rate expectations to solve the identification problem within GADTSMs, the OIS-augmented model should provide estimates of interest rate expectations that vary less with respect to the sample period in comparison to unaugmented models.

To assess the usefulness of the OIS-augmented GADTSM for real-time policy analysis, I compare real-time estimates of fitted and risk-neutral yields to estimates from the same model estimated using three different samples. To focus on monetary policy, I compare estimates on 19 different dates, the first 16 of which were major US unconventional monetary policy announcement days (see chapter 4, table 3).⁴ Chapter 4 demonstrates that many of these announcements significantly affected bond yields. The remaining 3 dates are: 30/04/2013, the end of the month prior to the May 2013 ‘taper tantrum’; 31/12/2015, the end of the term structure sample in chapter 4; and 31/12/2016.

I attain real-time estimates by re-estimating each model using data up to the date of interest. The start date of all real-time samples is January 2002. The three alternative samples all begin

⁴The 16 dates are: 25/11/2008; 01/12/2008; 16/12/2008; 28/01/2009; 18/03/2009; 10/08/2010; 27/08/2010; 21/09/2010; 15/10/2010; 03/11/2010; 09/08/2011; 26/08/2011; 25/01/2012; 31/08/2012; 13/09/2012; and 12/12/2012.

in January 2002 and end in April 2013, December 2015 and December 2016, respectively.

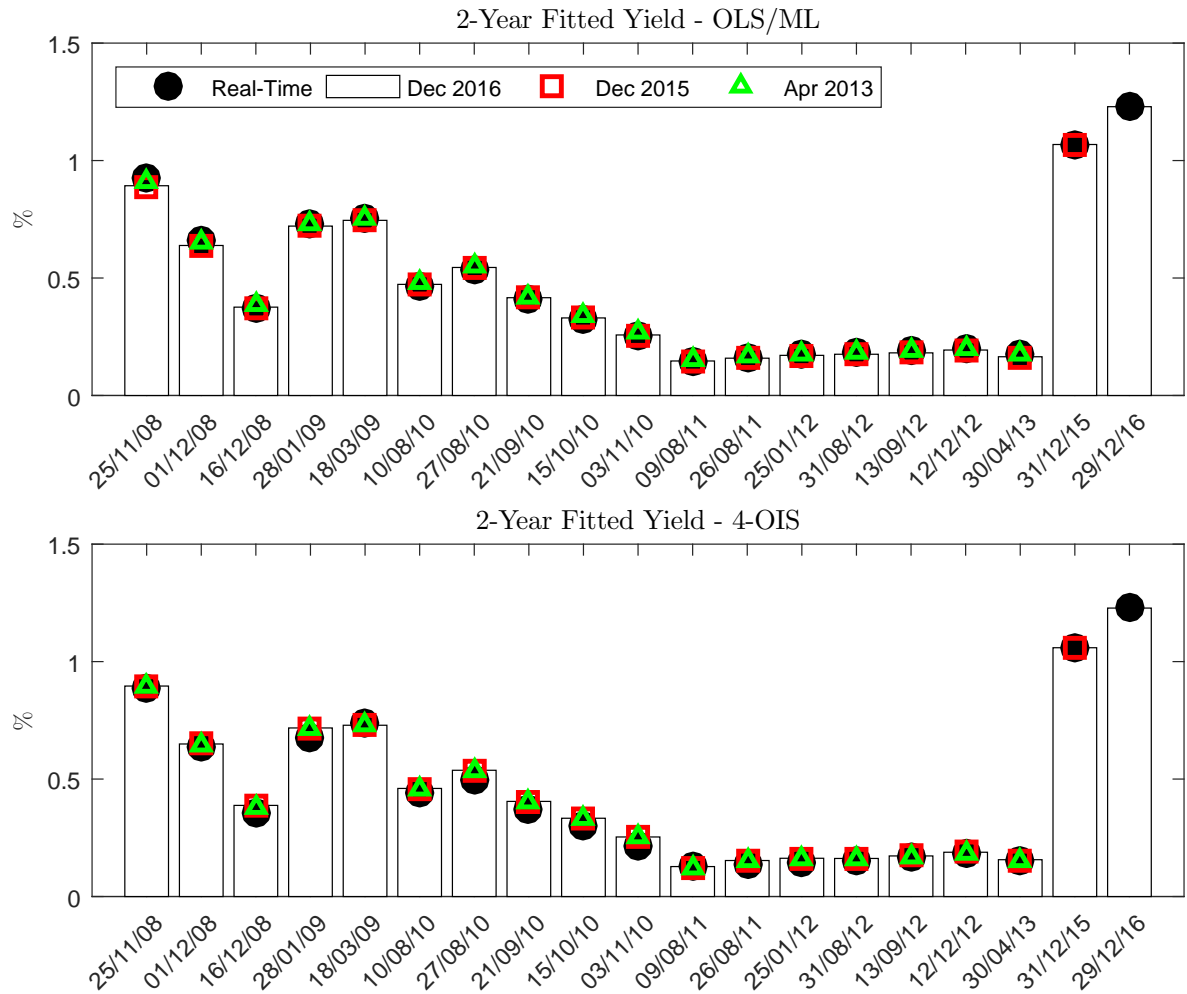
Figures B.1 and B.2 plot estimates of 2-year fitted yields and 2-year risk-neutral yields, respectively, from the unaugmented OLS/ML (top panel) and 4-OIS-augmented (bottom panel) models on the 19 different dates. The black dots represent the real-time estimates, which are most relevant for monetary policy. The white bars illustrate the 2-year fitted yields and 2-year risk-neutral yields from the January 2002 to December 2016 sample. The red squares plot the same quantities from the January 2002 to December 2015 sample, while the green triangles plot the results from the January 2002 to April 2013 sample. To attain reliable inference from an event study using a GADTSM, estimates of interest rate expectations and term premia on a given date from a desirable model should not vary across sample periods. For example, the estimated influence of the initial announcement of large-scale asset purchases (25/11/2008) on interest rate expectations and term premia should not change significantly as the estimation sample period is extended.

Figure B.1 illustrates that real-time estimates of the fitted yield are very similar to those attained using the three longer samples. This is unsurprising, as the identification problem pertains to the risk-neutral yields. However, the top panel of figure B.2 demonstrates that the identification problem does generate instability in risk-neutral yield estimates in the unaugmented OLS/ML model. Real-time estimates of the level of interest rate expectations differ substantially from those attained from the three longer samples. Moreover, the estimates from the three longer samples substantially differ from one another. For example, on March 18, 2009, the real-time estimate of the 2-year risk-neutral yield from the OLS/ML model is 111 basis points above the estimate attained from the January 2002 to December 2016 sample, which, in turn, is 29 basis points below the estimate from the January 2002 to April 2013 sample. In contrast, the bottom panel of figure B.2 illustrates that estimates of risk-neutral yields from the 4-OIS-augmented model are remarkably stable across samples. Although, there are differences between the real-time estimates and longer-sample estimates for early events that peak at 17 basis points on December 1, 2008 and December 16, 2008, this is likely to be due to parameter instability around 2008-2010. As the sample is extended to include more post-2008 data, the differences between estimates decline. On December 12, 2012, the range of estimates from the 4-OIS-augmented model is just 7 basis points; the corresponding figure for the OLS/ML model is 65 basis points. Moreover, the differences between real-time and longer-sample estimates for early events cannot be explained by small-sample issues, because shorter-sample estimates are close to real-time and longer-sample estimates on other event days. For instance, on event day 6 (August 10, 2010), the real-time estimate of the OIS-augmented 2-year yield is 48 basis points, the estimate from the 2002-2015 sample is 41 basis points, while an estimate using a 6.5-year sample from January 2004 to the event date is 45 basis points. Therefore, inference about interest rate expectations can be reliably made from the OIS-augmented model, regardless of the sample period chosen.

B.4.2 Four-Factor Specification

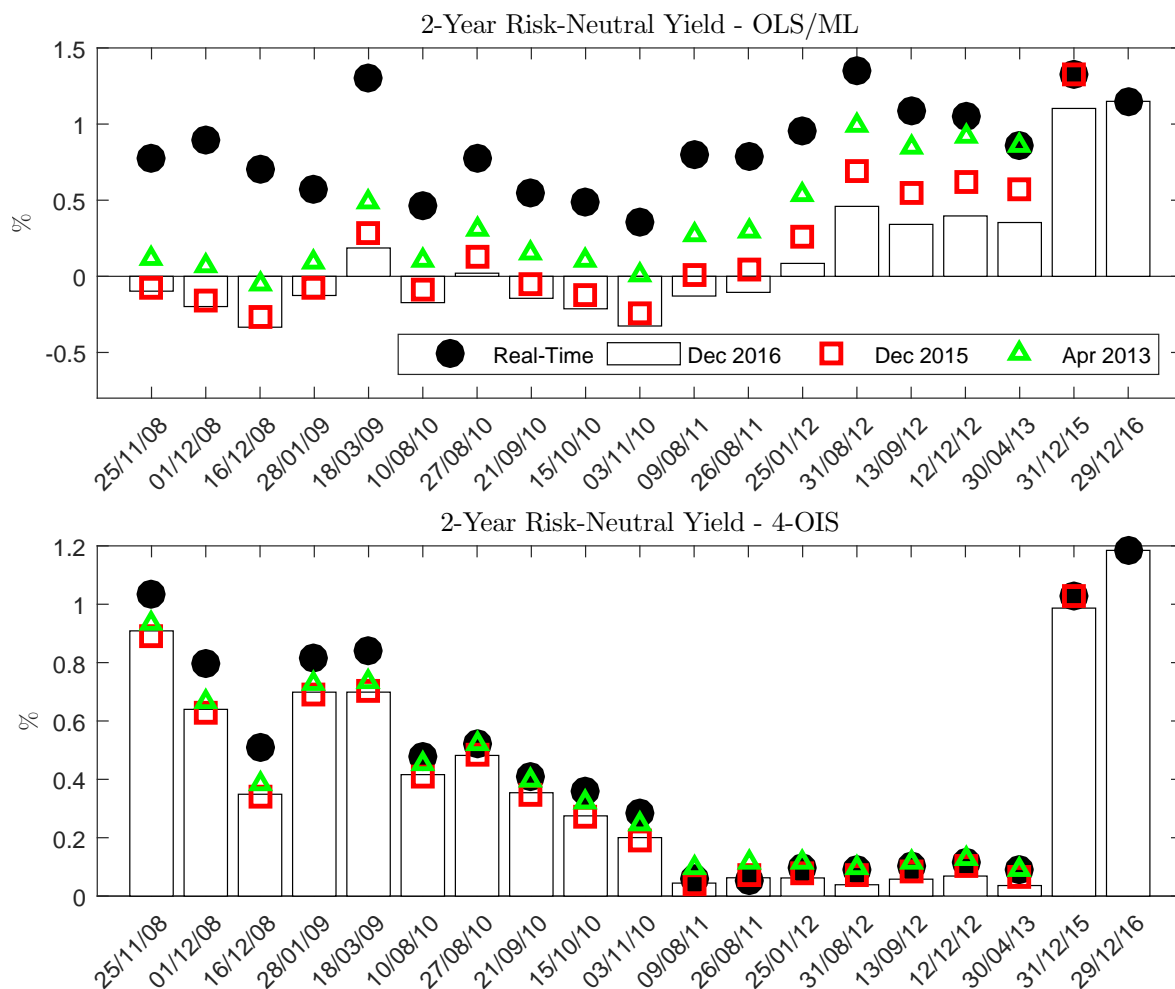
In the light of evidence by Cochrane and Piazzesi (2005, 2008) and Duffee (2011), who argue that more than three factors are necessary to explain the evolution of nominal Treasury yields, I estimate a four-factor specification of the GADTSMs. Although the four-factor model better fits actual bond yields for the 2002-2016 sample, I do not present these results in chapter 3 because

Figure B.1: Estimates of the 2-Year Fitted Yield from the Unaugmented OLS/ML and 4-OIS-Augmented GADTSMs on 19 Different Dates



Note: Estimates of the 2-year fitted yield from the unaugmented OLS/ML (top panel) and 4-OIS-augmented (bottom panel) models on 19 dates, 16 of which are associated with US unconventional monetary policy announcements in chapter 4. All samples begin in January 2002. The black dots represent real-time estimates of bond yields, using data up to the event date. The white bars represent estimates using a sample that ends in December 2016, the red squares represent estimates using a sample that ends in December 2015, and the green triangles represent estimates using a sample that ends in April 2013. All models are estimated with three pricing factors, using daily data. Yields are plotted in annualised percentage points. The date format for events, on the horizontal axis, is DD/MM/YY.

Figure B.2: Estimates of the 2-Year Risk-Neutral Yield from the Unaugmented OLS/ML and 4-OIS-Augmented GADTSMs on 19 Different Dates



Note: Estimates of the 2-year risk-neutral yield from the unaugmented OLS/ML (top panel) and 4-OIS-augmented (bottom panel) models on 19 dates, 16 of which are associated with US unconventional monetary policy announcements in chapter 4. All samples begin in January 2002. The black dots represent real-time estimates of risk-neutral yields, using data up to the event date. The white bars represent estimates using a sample that ends in December 2016, the red squares represent estimates using a sample that ends in December 2015, and the green triangles represent estimates using a sample that ends in April 2013. All models are estimated with three pricing factors, using daily data. Yields are plotted in annualised percentage points. The date format for events, on the horizontal axis, is DD/MM/YY.

the economic meaning of the pricing factors in a three-factor model is well understood (i.e., level, slope and curvature), while the economic interpretation the fourth factor is less well understood. Moreover, the differences in risk-neutral yield estimates from the three and four-factor models are small, and there is no evidence that a single model is unambiguously preferable. Lloyd (2017a, Appendix F.2) provides further details, indicating that: (i) the fitted yields from the four-factor GADTSMs do not differ markedly from one another; (ii) as with the three-factor models, the four-factor OIS-augmented models accurately fit OIS rates, but the three-factor OIS-augmented models perform marginally better than their four-factor counterparts on a RMSE basis; (iii) the inclusion of OIS rates in the estimation of four-factor GADTSMs does not significantly influence the bond pricing factors; and (iv) the risk-neutral forward yields from the four-factor OIS-augmented models are similar to those from their three-factor variants plotted in figure 3.6, but a comparison of the four and three-factor models indicates that neither unambiguously outperforms the other when compared to corresponding-horizon federal funds futures rates.

Appendix C

Appendix to Chapter 4

C.1 Event Study

C.1.1 Event-Specific Explanations

Table C.1: Detailed Explanation of Increase in Bond Yields on Event Days IV, VII, IX and XVI

#	Date	Explanation
IV	28/01/2009	On this date, the Fed announced continued asset purchases and a readiness to expand the programme if conditions warranted. The increase in bond yields and OIS rates on this date can be rationalised when considering the pre-announcement expectations of market participants. Anecdotal evidence indicates that market participants were disappointed by the lack of concrete language regarding the possibility and timing of purchases of longer-dated Treasury securities. Sources: Neely (2010); Bauer and Rudebusch (2014).
VII	27/08/2010	On this date, Chairman Bernanke spoke at Jackson Hole, stating that “additional purchases [...] would be effective.” According to Reuters, “[T]he overall tone [of the speech] was one of watch and wait,” Goldman Sachs economist Jan Hatzius wrote in a note to clients. [...] Stocks initially fell after Bernanke’s remarks, but reversed course and three major indexes closed up 1.7 percent.” On the same day, the US Commerce Department cut its estimate for US GDP in 2010 Q2 from 2.4% to 1.6%, but this was higher than the 1.4% surveys had predicted. Source: in.reuters.com/article/columns-us-usa-fed-bernanke-idINTRE6700MF20100828 .
IX	15/10/2010	On this date, Chairman Bernanke said the central bank has “a case for further action” to stimulate the economy, citing high unemployment and low inflation. However, markets did not learn specifics on future LSAPs from the speech. “We’re still no closer to knowing exactly what the so-called QE2 program might involve, in particular what amount of Treasury securities the Fed could end up buying,” said Paul Ashworth, an economist at Capital Economics. Although stocks opened higher after Chairman Bernanke’s speech, the tone cooled as investors received other economic news, including US sales growth of 0.6% in September 2010, which was higher than anticipated. Source: money.cnn.com/2010/10/15/markets/markets_newyork/index.htm .

Table C.1: Detailed Explanation of Increase in Bond Yields on Event Days IV, VII, IX and XVI

#	Date	Explanation
XVI	12/12/2012	On this date, the Fed announced state-dependent threshold-based forward guidance. Contemporary news reports stated that this was “aggressively dovish”, indicating that the policy was perceived to be expansionary. However, on the same day, the FOMC downgraded its forecast for the US economy in 2013 2.5-3.0% year-on-year to 2.3-3.0%. “Stock prices jumped after the Fed released its policy statement at midday, then began falling during Mr Bernanke’s news conference about two hours later as he insisted the Fed was not significantly increasing its efforts to bolster the economy.” Source: www.bloomberg.com/news/articles/2012-12-12/the-fed-turns-aggressively-dovish-with-evans-rule and www.nytimes.com/2012/12/13/business/economy/fed-to-maintain-stimulus-bond-buying.html?mcubz=1 .

C.1.2 Event Significance

To assess the significance of daily changes in actual yields yld , OIS rates i^{ois} , fitted yields \widehat{yld} , risk-neutral yields \widehat{exp} and term premia \widehat{tp} at the 2, 5 and 10-year maturities on specific event days (considered in tables 4.4-4.7), I estimate regressions of the following form:

$$\Delta x_{n,t} = \alpha_{x,n} + \mathbf{Event}_t \beta_{x,n} + \mathbf{D}_t \gamma_{x,n} + \varepsilon_{x,n,t} \quad (\text{C.1})$$

where $x = \{yld, i^{ois}, \widehat{yld}, \widehat{exp}, \widehat{tp}\}$ and $n = \{2, 5, 10\}$. \mathbf{Event}_t is a 1×16 vector of 16 dummy variables each pertaining to a specific unconventional monetary policy announcement date listed in table 4.3. The dummies are set equal to 1 on the announcement date they are linked with, and 0 otherwise. $\beta_{x,n}$ is a 16×1 vector of parameters to be estimated. The control variables are included in \mathbf{D}_t , a $1 \times K$ vector of K dummy variables pertaining to other macroeconomic data releases, set equal to 1 on release dates and 0 otherwise, where K is the number of different types of macroeconomic data releases (12). $\gamma_{x,n}$ is a $K \times 1$ vector of parameters to be estimated. The twelve macroeconomic data releases, from the US Bureau of Economic Analysis and Bureau of Labour Statistics, included are: Advanced US GDP; Preliminary US GDP; Final US GDP; International Trade; US Personal Income and Outlays; US International Transactions; US Employment Situation; US Job Openings and Labour Turnover; US Producer Price Index; US Consumer Price Index; Preliminary US Productivity and Costs; and Revised US Productivity and Costs.

The i -th element of $\beta_{x,n}$ represents the difference between the change in x on unconventional monetary policy announcement day i , where $i = 1, 2, \dots, 16$, and the average daily change in x on other dates, excluding other unconventional monetary policy announcements and other US macroeconomic data releases. If the i -th element of $\widehat{\beta}_{x,n}$ is statistically significant, then unconventional policy announcement i is said to have a significant effect on x at the n -year horizon. The additional macroeconomic data release dummies \mathbf{D}_t are included to ensure that the significance of daily changes in x are compared to dates with no significant news pertaining to the US economy. This empirical specification underpins the significance levels associated with daily changes in variables in tables 4.4-4.7.

Table C.2 presents the significance of all 16 event dates together. To assess the significance

Table C.2: Significance of Daily Changes in US Treasury Yields and Their Sub-Components from Equation (C.2)

Model	$\widehat{\beta}_{yld,n}$	$\widehat{\beta}_{exp,n}$	$\widehat{\beta}_{tp,n}$
2-Year Maturity			
Bias-Corrected	−5.06***	−2.85***	−2.21
Survey-Augmented	−4.53***	−0.84*	−3.70**
OIS-Augmented	−4.65***	−4.36***	−0.29***
5-Year Maturity			
Bias-Corrected	−7.73**	−3.17**	−4.56
Survey-Augmented	−7.86**	−2.29***	−5.57**
OIS-Augmented	−6.96***	−6.26***	−0.70*
10-Year Maturity			
Bias-Corrected	−8.00*	−2.75	−5.26
Survey-Augmented	−9.28**	−2.69***	−6.59**
OIS-Augmented	−6.51**	−4.84***	−1.67

Note: Estimates of $\beta_{x,n}$ from (C.2) using daily frequency changes in fitted yields, risk-neutral yields and term premia from the bias-corrected, survey-augmented and OIS-augmented GADTSMs using data from January 2, 2008 to April 31, 2013. Estimates that are significant at the 1%, 5% and 10% significance level are denoted with asterisks ***, ** and * respectively. Significance is determined by t -statistics calculated using Newey and West (1987) standard errors. All figures are reported in basis points to two decimal places.

of daily changes in fitted yields \widehat{yld} , risk-neutral yields \widehat{exp} and term premia \widehat{tp} at the 2, 5 and 10-year maturities on all event days, I estimate regressions of the following form:

$$\Delta x_{n,t} = \alpha_{x,n} + \beta_{x,n} event_t + \mathbf{D}_t \gamma_{x,n} + \varepsilon_{x,n,t} \quad (\text{C.2})$$

$event_t$ is a dummy variable equal to 1 on the LSAP and forward guidance event dates in table 4.3, and 0 otherwise. $\beta_{x,n}$ is a scalar to be estimated.

$\beta_{x,n}$ represents the difference between average daily changes in x on unconventional monetary policy announcement days and on other dates that exclude other US macroeconomic data releases. If $\widehat{\beta}_{x,n}$ is statistically significant, then the 16 unconventional monetary policy announcements are said to have a cumulative significant effect on x at the n -year horizon.

Table C.2 presents the results of estimation of (C.2). The results indicate that the daily changes of fitted yields from all three models at all three maturities on unconventional monetary policy announcement days were significantly different to their daily change on other dates, excluding other macroeconomic data release dates. Moreover, changes in interest rate expectations from the OIS-augmented decomposition were also significantly different on unconventional monetary policy announcement dates at all horizons.

Appendix D

Appendix to Chapter 5

D.1 Model Derivation

This appendix provides mathematical derivations for expressions in section 5.3.

D.1.1 Households

The Home household problem maximises its discounted expected lifetime utility (5.1) subject to (5.8), the Home household budget constraint in units of the Home consumption basket. The associated optimality conditions are given by (5.12)-(5.13), where $r_t = R_{t-1}/\pi_t$, $r_t^* = Q_t/(Q_{t-1}P_t^*Z_{t-1}^*)$, $r_{L,t} = (R_{L,t}P_{L,t})/(\pi_t P_{L,t-1})$ and $r_{L,t}^* = (R_{L,t}^*\mathcal{E}_t P_{L,t}^*)/(\pi_t \mathcal{E}_{t-1} P_{L,t-1}^*)$.

The Home household budget constraint can be expressed in nominal terms:

$$\begin{aligned} P_t C_t + \check{B}_{HH,t} + e^{-\zeta_{F,t}} \mathcal{E}_t \check{B}_{HF,t} + P_{L,t} \check{B}_{HH,t}^L + e^{-\zeta_{F,t}} \mathcal{E}_t P_{L,t}^* \check{B}_{HF,t}^L + e^{-\zeta_{S,t}} \check{S}_{HH,t} + \\ e^{-\zeta_{S,t}^*} e^{-\zeta_{F,t}} \mathcal{E}_t \check{S}_{HF,t} = P_t w_t L_t + P_t \Pi_t - P_t T_t + \check{B}_{HH,t-1} R_{t-1} + \check{B}_{HF,t-1} \mathcal{E}_t R_{t-1}^* + \\ \check{B}_{HH,t-1}^L (1 + \kappa P_{L,t}) + \check{B}_{HF,t-1}^L \mathcal{E}_t (1 + \kappa P_{L,t}^*) + \check{S}_{HH,t-1} \frac{P_t}{P_{t-1}} r_{e,t} + \check{S}_{HF,t-1} \frac{P_t \mathcal{E}_{t-1}}{P_{t-1}} r_{e,t}^* \end{aligned}$$

where $\check{B}_{jk,t}$, $\check{B}_{jk,t}^L$ and $\check{S}_{jk,t}$ denote country- j holdings of country- k short-term bonds, long-term bonds and equity, respectively, in units of issuing country's currency. To attain the Home household budget constraint in units of the Home consumption basket (5.8) from this, divide through by P_t , define

$$\begin{aligned} B_{HH,t} = \frac{\check{B}_{HH,t}}{P_t}, \quad B_{HF,t} = \frac{\mathcal{E}_t \check{B}_{HF,t}}{P_t}, \quad B_{HH,t}^L = \frac{P_{L,t} \check{B}_{HH,t}^L}{P_t}, \\ B_{HF,t}^L = \frac{\mathcal{E}_t P_{L,t}^* \check{B}_{HF,t}^L}{P_t}, \quad S_{HH,t} = \frac{\check{S}_{HH,t}}{P_t}, \quad S_{HF,t} = \frac{\mathcal{E}_t \check{S}_{HF,t}}{P_t} \end{aligned}$$

and use the definitions of r_t , r_t^* , $r_{L,t}$, $r_{L,t}^*$, $r_{e,t}$ and $r_{e,t}^*$ provided in section 5.3.1.

The Foreign household problem maximises its discounted expected lifetime utility subject to the Foreign household budget constraint, expressed in units of the Foreign consumption basket:

$$\begin{aligned} C_t^* + Q_t^{-1} \left(B_{FH,t} + e^{-\zeta_{F,t}} B_{FF,t} + B_{FH,t}^L + e^{-\zeta_{F,t}} B_{FF,t}^L + e^{-\zeta_{S,t}} S_{FH,t} + e^{-\zeta_{F,t}} e^{-\zeta_{S,t}^*} S_{FF,t} \right) \\ = w_t^* L_t^* + \Pi_t^* - T_t^* + Q_t^{-1} \left(B_{FH,t-1} r_t + B_{FF,t-1} r_t^* + B_{FH,t-1}^L r_{L,t} + \right. \\ \left. B_{FF,t-1}^L r_{L,t}^* + S_{FH,t-1} r_{e,t} + S_{FF,t-1} r_{e,t}^* \right) \end{aligned}$$

The associated optimality conditions are:

$$w_t^* = e^{\zeta_{L,t}^*} (C_{X,t}^*)^\sigma (L_t^*)^{\frac{1}{\nu}} \quad (\text{D.1})$$

$$1 = \mathbb{E}_t \left[\beta (C_t^*) e^{\Delta \zeta_{C,t+1}^*} e^{\zeta_{i,t}} \left(\frac{C_{X,t+1}^*}{C_{X,t}^*} \right)^{-\sigma} \frac{Q_t}{Q_{t+1}} r_{i,t+1} \right] \quad \text{where } i = 1, 2, \dots, 6 \quad (\text{D.2})$$

where $C_{X,t}^* \equiv C_t^* - \gamma C_{t-1}^*$, $\zeta_{1,t} \equiv 0$, $\zeta_{2,t} \equiv \zeta_{F,t}$, $\zeta_{3,t} \equiv 0$, $\zeta_{4,t} \equiv \zeta_{F,t}$, $\zeta_{5,t} \equiv \zeta_{S,t}$, $\zeta_{6,t} \equiv \zeta_{F,t} + \zeta_{S^*,t}$, $r_{1,t+1} \equiv r_{t+1}$, $r_{2,t+1} \equiv r_{t+1}^*$, $r_{3,t+1} \equiv r_{L,t+1}$, $r_{4,t+1} \equiv r_{L,t+1}^*$, $r_{5,t+1} \equiv r_{e,t}$, and $r_{6,t+1} \equiv r_{e,t}^*$.

D.1.2 Firms

Output in each country is produced by a continuum of monopolistically competitive firms. The production function for a good produced by firm $h \in (0, 1)$ in the Home country is (5.14).

Marginal Cost of Production

The Home firm chooses factor inputs to minimise real total costs tc_t :

$$\min_{\{L_t(h), X_t(h)\}} tc_t(h) = w_t L_t(h) + X_t(h)$$

subject to their production function (5.14). This yields the following optimal factor demands:

$$L_t(h) = Y_t(h) e^{-a_t} \left(\frac{1}{w_t} \frac{1-\alpha}{\alpha} \right)^\alpha, \quad X_t(h) = Y_t(h) e^{-a_t} \left(w_t \frac{\alpha}{1-\alpha} \right)^{1-\alpha}$$

With these factor demands, real total costs for firm h are:

$$tc_t(h) = Y_t(h) e^{-a_t} \frac{w_t^{1-\alpha}}{(1-\alpha)^{1-\alpha} \alpha^\alpha}$$

so the real marginal cost function for the Home firm is:

$$mc_t = e^{-a_t} \frac{w_t^{1-\alpha}}{(1-\alpha)^{1-\alpha} \alpha^\alpha}$$

which is the same for all Home firms, because each firm faces the same aggregate productivity shock and real wage. A similar expression exists for Foreign firms:

$$mc_t^* = e^{-a_t^*} \frac{(w_t^*)^{1-\alpha}}{(1-\alpha)^{1-\alpha} \alpha^\alpha} \quad (\text{D.3})$$

The optimal input demand functions for the Home firms, written in terms of the real marginal cost, are (5.17) and (5.18). Similar expressions are attained for the Foreign firm:

$$L_t^* = \frac{(1-\alpha) mc_t^* Y_t^*}{w_t^*} \quad (\text{D.4})$$

$$X_t^* = \alpha mc_t^* Y_t^* \quad (\text{D.5})$$

Pricing: Producer Currency Pricing (PCP)

Under PCP, the Home firm's pricing problem is given by (5.24), with associated optimality conditions (5.25).

The Home good price index is given by:

$$P_{H,t} = \left[\int_0^1 P_t(h)^{1-\theta} dh \right]^{\frac{1}{1-\theta}}$$

This comprises surviving contracts and newly set prices. Given that in each period a price contract has a probability $1 - \xi$ of ending, the probability that a contract signed in period $t - s$ survives until period t and ends in that period is $(1 - \xi)\xi^s$. Therefore:

$$P_{H,t} = \left[\int_0^1 \sum_{s=0}^{\infty} (1 - \xi)\xi^s \mathcal{P}_{t-s}(h)^{1-\theta} dh \right]^{\frac{1}{1-\theta}}$$

so with PCP, the price indices evolve according to (5.27) and (5.28).

D.1.3 Equilibrium

Here, I derive (5.34). First, set all shocks to zero in (5.8) and impose the government budget constraint, $T_t = G_t \equiv 0$:

$$C_t + B_{HH,t} + B_{HF,t} + B_{HH,t}^L + B_{HF,t}^L + S_{HH,t} + S_{HF,t} = w_t L_t + \Pi_t + B_{HH,t-1} r_t + B_{HF,t-1} r_t^* + B_{HH,t-1}^L r_{L,t} + B_{HF,t-1}^L r_{L,t}^* + S_{HH,t-1} r_{e,t} + S_{HF,t-1} r_{e,t}^*$$

Then, substitute for real profits Π_t using (5.19):

$$\begin{aligned} P_t C_t + P_t (B_{HH,t} + B_{HF,t} + B_{HH,t}^L + B_{HF,t}^L + S_{HH,t} + S_{HF,t}) = \\ P_t w_t L_t + P_{H,t} Y_{H,t} + \mathcal{E}_t P_{H,t}^* Y_{H,t}^* - P_t m c_t (Y_{H,t} + Y_{H,t}^*) + \\ P_t (B_{HH,t-1} r_t + B_{HF,t-1} r_t^* + B_{HH,t-1}^L r_{L,t} + B_{HF,t-1}^L r_{L,t}^* + S_{HH,t-1} r_{e,t} + S_{HF,t-1} r_{e,t}^*) \end{aligned}$$

which can be rewritten as:

$$P_t C_t + P_t \sum_{i=1}^6 \alpha_{i,t} = P_t (1 - \alpha) m c_t Y_t + P_{H,t} Y_{H,t} + \mathcal{E}_t P_{H,t}^* Y_{H,t}^* - P_t m c_t Y_t + P_t \sum_{i=1}^6 \alpha_{i,t-1} r_{i,t}$$

using (5.17) and (5.31), where $\alpha_{i,t}$ and $r_{i,t}$ are defined in the main body of the text. Rearranging:

$$P_t (C_t + \alpha m c_t Y_t) - P_{H,t} Y_{H,t} + P_t \sum_{i=1}^6 \alpha_{i,t} - P_t \sum_{i=1}^6 \alpha_{i,t-1} r_{i,t} = \mathcal{E}_t P_{H,t}^* Y_{H,t}^*$$

Using (5.20) and (5.21), in a symmetric steady state:

$$P_t \sum_{i=1}^6 \alpha_{i,t} - P_t \sum_{i=1}^6 \alpha_{i,t-1} r_{i,t} = \mathcal{E}_t P_{H,t}^* Y_{H,t}^* - P_{F,t} Y_{F,t}$$

returning (5.34).

D.2 Alternative Price-Setting Regimes

In this appendix, I present two alternative price-setting regimes to PCP: local currency pricing (LCP) (Betts and Devereux, 2000), where firms set prices in the currency of the market in which they sell the goods; and dollar currency pricing (DCP) (Gopinath, 2015), where both the home country and the US (the foreign country) invoice their export prices in dollar terms.

Local Currency Pricing As with PCP, let $\mathcal{P}_t(h)$ denote the price of the Home good in the Home market optimally chosen by firm h who can reset their price at time t . However, unlike under PCP, the Home firm h sets its Foreign price in Foreign currency terms $\mathcal{P}_t^*(h)$ under LCP. The pricing problem for Home firms under LCP can be written as:

$$\max_{\{\mathcal{P}_t(h), \mathcal{P}_t^*(h)\}} \mathbb{E}_t \sum_{s=0}^{\infty} \delta_s \xi^s \Omega_{t+s} \left[\frac{(\mathcal{P}_t(h))^{1-\theta}}{(P_{H,t+s})^{-\theta}} Y_{H,t+s} + \mathcal{E}_{t+s} \frac{(\mathcal{P}_t^*(h))^{1-\theta}}{(P_{H,t+s}^*)^{-\theta}} Y_{H,t+s}^* - P_{t+s} m c_{t+s} \left(\frac{\mathcal{P}_t(h)}{P_{H,t+s}} \right)^{-\theta} Y_{H,t+s} - P_{t+s} m c_{t+s} \left(\frac{\mathcal{P}_t^*(h)}{P_{H,t+s}^*} \right)^{-\theta} Y_{H,t+s}^* \right]$$

The Home firms' optimality conditions are:

$$\begin{aligned} \mathcal{P}_t(h) &= e^{\mu_t} \frac{\theta}{\theta - 1} \frac{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s \xi^s \Omega_{t+s} m c_{t+s} P_{t+s} P_{H,t+s}^{\theta} Y_{H,t+s}}{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s \xi^s \Omega_{t+s} P_{H,t+s}^{\theta} Y_{H,t+s}} \\ \mathcal{P}_t^*(h) &= e^{\mu_t} \frac{\theta}{\theta - 1} \frac{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s \xi^s \Omega_{t+s} m c_{t+s} P_{t+s} (P_{H,t+s}^*)^{\theta} Y_{H,t+s}^*}{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s \xi^s \Omega_{t+s} \mathcal{E}_{t+s} (P_{H,t+s}^*)^{\theta} Y_{H,t+s}^*} \end{aligned} \quad (\text{D.6})$$

where the markup shock μ_t is defined in (5.26).

The Foreign optimality conditions are:

$$\begin{aligned} \mathcal{P}_t(f) &= e^{\mu_t^*} \frac{\theta}{\theta - 1} \frac{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s^* \xi^s \Omega_{t+s}^* m c_{t+s}^* P_{t+s}^* P_{F,t+s}^{\theta} Y_{F,t+s}}{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s^* \xi^s \Omega_{t+s}^* \mathcal{E}_{t+s}^{-1} P_{F,t+s}^{\theta} Y_{F,t+s}} \\ \mathcal{P}_t^*(f) &= e^{\mu_t^*} \frac{\theta}{\theta - 1} \frac{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s^* \xi^s \Omega_{t+s}^* m c_{t+s}^* P_{t+s}^* (P_{F,t+s}^*)^{\theta} Y_{F,t+s}^*}{\mathbb{E}_t \sum_{s=0}^{\infty} \delta_s^* \xi^s \Omega_{t+s}^* (P_{F,t+s}^*)^{\theta} Y_{F,t+s}^*} \end{aligned} \quad (\text{D.7})$$

where $\Omega_{t+s}^* \equiv \frac{u'(C_{t+s}^*)}{u'(C_t^*)}$ is the discount factor used to evaluate Foreign firm profits.

Because Home and Foreign prices are set independently, the law of one price is violated with any movement of the exchange rate under LCP. Since all producers that reset their price in period t will choose the same price level, there are now four equations that describe the dynamic evolution of the price indices $P_{H,t}$, $P_{H,t}^*$, $P_{F,t}$ and $P_{F,t}^*$:

$$P_{H,t}^{1-\theta} = \xi P_{H,t-1}^{1-\theta} + (1 - \xi) \mathcal{P}_t(h)^{1-\theta} \quad (\text{D.8})$$

$$(P_{H,t}^*)^{1-\theta} = \xi (P_{H,t-1}^*)^{1-\theta} + (1 - \xi) \mathcal{P}_t^*(h)^{1-\theta} \quad (\text{D.9})$$

$$P_{F,t}^{1-\theta} = \xi P_{F,t-1}^{1-\theta} + (1 - \xi) \mathcal{P}_t(f)^{1-\theta} \quad (\text{D.10})$$

$$(P_{F,t}^*)^{1-\theta} = \xi (P_{F,t-1}^*)^{1-\theta} + (1 - \xi) \mathcal{P}_t^*(f)^{1-\theta} \quad (\text{D.11})$$

Dollar Currency Pricing Gopinath (2015) documents that an overwhelming share of international trade is invoiced in very few currencies, with the dollar the dominant currency. Moti-

vated by this, I investigate the effect of DCP on international risk sharing and the international transmission of shocks.

To implement DCP, firms in the Home country act isomorphically to LCP; they set the domestic prices in the Home currency and their export prices in the Foreign currency — here, the US dollar. In contrast, Foreign firms engage in PCP; they set both domestic and export prices in the Foreign currency.

Matching the Stylised Facts Under Different Price-Setting Regimes Table D.1 illustrates that the model differs in its ability to match the empirical evidence under different price-setting regimes. Here, I present the steady state portfolio quantities for Home investors under the model’s baseline calibration presented in table 5.2. The PCP model most closely matches the stylised facts at the baseline trade elasticity ϕ parameterisation of 0.78. Under LCP and DCP, the model quantitatively replicates equity home bias \tilde{s}_{HH} , Home holdings of US-issued equity s_{HF} , and Home holdings of Home-issued short-term debt b_{HH}^S . Both LCP and DCP models are unable to qualitatively match empirical evidence related to holdings of long-term debt and total debt.

Table D.2 presents the ranges of value of ϕ within which the steady state portfolio quantities for Home investors are consistent with the empirical evidence.

Table D.1: Steady State Portfolio Quantities for Home Investors Under the Model's Baseline Calibration

HOLDINGS	DATA	PCP	LCP	DCP
Total US Assets, t_{HF}	5.54-33.66%	5.68%	-2.81%	<i>1.96%</i>
Equity Home Bias, \tilde{s}_{HH}	50-100%	97.09%	93.29%	95.14%
US Equity, s_{HF}	2.70-20.71%	2.77%	6.46%	4.66%
Home Debt, b_{HH}	<0%	- 3.71%	7.60%	1.32%
US Debt, b_{HF}	2.84-12.95%	2.90%	-9.26%	-2.70
Home Short-Term Debt, b_{HH}^S	<0%	- 2.79%	- 4.80%	- 4.45%
US Short-Term Debt, b_{HF}^S	0.28-1.25%	<i>1.97%</i>	<i>3.12%</i>	<i>3.08%</i>
Home Long-Term Debt, b_{HH}^L	<0%	- 0.93%	12.40%	5.77%
US Long-Term Debt, b_{HF}^L	2.20-12.17%	<i>0.93%</i>	-12.39%	-5.78%

Notes: Steady state portfolio quantities under the baseline model calibration, including $\phi = 0.78$. Emboldened values denote quantities that quantitatively match the empirical evidence. Italicised values denote quantities that qualitatively match the empirical evidence. Values that are neither emboldened nor italicised do not match the empirical evidence qualitatively or quantitatively.

Table D.2: Ranges of Values of the Trade Elasticity ϕ within which Steady State Portfolio Quantities for Home Investors are Consistent with Empirical Evidence

HOLDINGS	PCP	LCP	DCP
Total US Assets, t_{HF}	0.690-0.785	0.660-0.755	0.675-0.770
Equity Home Bias, \tilde{s}_{HH}	0.185-0.205, 0.640-0.790	0.185-0.200, 0.650-0.805	0.185-0.200, 0.645-0.795
US Equity, s_{HF}	0.725-0.785	0.735-0.795	0.730-0.790
Home Debt, b_{HH}	0.005-0.880	0.005-0.195, 0.205-0.485	0.005-0.740
US Debt, b_{HF}	0.325-0.380, 0.775-1.295	0.275-0.290, 1.045-1.345	0.295-0.320, 0.940-1.320
Home Short-Term Debt, b_{HH}^S	0.005-0.790	0.005-0.790	0.010-0.790
US Short-Term Debt, b_{HF}^S	0.780-0.785	0.780-0.790	0.780-0.790
Home Long-Term Debt, b_{HH}^L	0.775-2.500	0.205-0.215, 0.805-2.500	0.790-2.500
US Long-Term Debt, b_{HF}^L	0.780-0.805	0.810-0.835	0.795-0.820

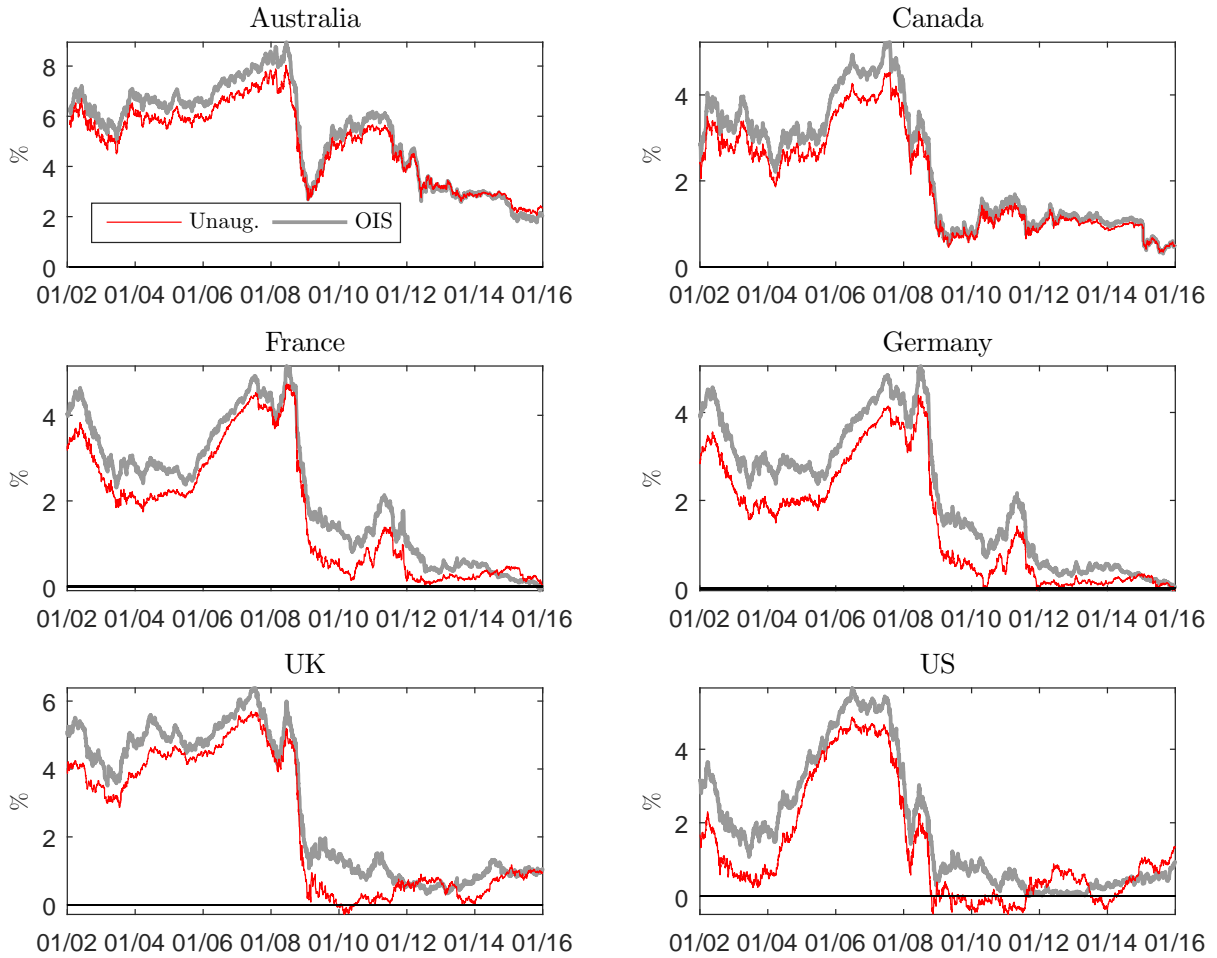
Notes: Ranges of values of the trade elasticity ϕ within which the steady state portfolio quantities of the model, presented in section 5.3, are consistent with the empirical evidence presented in section 5.2.2. I solve the model for values of ϕ ranging from 0.005 to 2.500, in increments of 0.005.

D.3 Estimates of the OIS-Augmented GADTSM

In this appendix, I present additional information regarding the estimates of the OIS-augmented GADTSM for Australia, Canada, France, Germany, the UK and the US. This is the first paper to apply the proposal of chapter 3 to countries other than the US. The results further support the conclusions in chapter 3.

Figure D.1 presents the 2-year risk-neutral interest rate expectations from the unaugmented and OIS-augmented models. Although the risk-neutral yields from the two models follow broadly similar patterns, risk-neutral yields from the unaugmented model are more often negative.

Figure D.1: 2-Year Risk-Neutral Yield from the Unaugmented and OIS-Augmented GADTSMs, January 2002 to December 2015



Note: 2-year risk-neutral yields from two GADTSMs: (i) unaugmented GADTSM, estimated using the algorithm of Joslin et al. (2011); and (ii) the OIS-augmented GADTSM, estimated using the algorithm of Lloyd (2017a). The models are estimated with three pricing factors, using daily data from January 2, 2002 to December 31, 2015. All figures are in annualised percentage points.

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